

Alternative formulas for synthetic dual system estimation in the 2000 census

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Abstract: The U.S. Census Bureau provides an estimate of the true population as a supplement to the basic census numbers. This estimate is constructed from data in a post-censal survey. The overall procedure is referred to as dual system estimation. Dual system estimation is designed to produce revised estimates at all levels of geography, via a synthetic estimation procedure.

We design three alternative formulas for dual system estimation and investigate the differences in area estimates produced as a result of using those formulas. The primary target of this exercise is to better understand the nature of the homogeneity assumptions involved in dual system estimation and their consequences when used for the enumeration data that occurs in an actual large scale application like the Census. (Assumptions of this nature are sometimes collectively referred to as the “synthetic assumption” for dual system estimation.)

The specific focus of our study is the treatment of the category of census counts referred to as imputations in dual system estimation. Our results show the degree to which varying treatment of these imputation counts can result in differences in population estimates for local areas such as states or counties.

1. Introduction

The U.S. census is required by the Constitution to be conducted every ten years. In an attempt to provide better estimates of the true population than contained in the basic census counts, the Census Bureau [13] uses both statistical and demographic methods. In 2000 the statistical process was called Accuracy and Coverage Evaluation (A.C.E.).

The 2000 A.C.E. data consists of two parts: the Population sample (P-sample) and the Enumeration sample (E-sample). The P-sample includes persons who are validly included in the A.C.E. survey, and the E-sample includes census enumerations from households in the A.C.E. block clusters. For a detailed overview of the 2000 A.C.E., please see Hogan [9] and Norwood and Citro [11].

The 2000 A.C.E. was designed to get an estimate of the population at every geographic level, based on the census count and the information from the E-sample and the P-sample. To be more precise, the procedure adopted by the Census Bureau is termed a synthetic dual system estimate. Its validity rests on several assumptions, including a major synthetic (homogeneity) assumption.

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Various technical assumptions can be made for synthetic assumption. These affect the details of the formulas needed to produce the final population estimates. For ideal and homogeneous populations any of the resulting formulas will produce unbiased estimates. However, the U.S. population does not appear to have this type of ideal structure. Hence different synthetic assumptions yield different estimates, and it does not appear that all of these estimates are actually unbiased.

This paper investigates the nature of these assumptions and the extent of the differences produced when using three alternative dual system formulas within the 2000 U.S. Census. It should be emphasized that the data available to us do not allow us to make any confident claim as to which of the estimates is more accurate; indeed such a claim is not our objective. Instead, we present our analyses as a means of providing better understanding of the dual system estimation process in the presence of actual populations, such as that encountered in the 2000 Census, and of judging the extent of differences that may be expected to result from differing assumptions about the census enumeration process.

Our analysis revolves around the extent and homogeneity of imputations of household and whole person records into the census enumeration. The available data allows us to produce alternative estimates based on different treatment of these imputations. As we later remark, there are other aspects of the dual system process that might involve analogous biases in the presence of inhomogeneity, however the data available to us do not allow for as complete an analysis relative to those factors.

In Section 2 we briefly discuss the nature and extent of imputation in the 2000 census. It is clear that the desired stochastic homogeneity does not hold there. Section 3 introduces background for dual system estimation and the synthetic assumption. The alternative formulas are presented in Section 4. Section 5 displays the results of using these formulas to estimate the true population shares of the states in 2000. Section 6 presents similar results for estimation of population shares of groups of counties. Mathematical comparison of different formulas is made in Section 7. Section 8 contains a summary conclusion and remarks.

The data for A.C.E. was collected during the 2000 census and first prepared and analyzed before April 2001. The Census Bureau decided not to issue the results then produced as official census estimates. Following this, the data was re-analyzed several times, leading up to revised A.C.E. estimates, referred to as A.C.E. Revision II. These were released on March 2003. The revised data identified, and deleted from the estimation process, a significant number of records that were judged to be duplicates. There were also a number of other more technical, but not insignificant, innovations in A.C.E. Revision II. See Kostanich [10] for a more complete description of A.C.E. Revision II.

The analyses of our paper are based on the original April 2001 A.C.E. data. There are several reasons for our using this original data, rather than the revised A.C.E. II data. The primary reason is that this is the data that was supplied to us by the Bureau, beginning in 2001. (We gratefully acknowledge the Bureau's assistance in supplying us with suitable versions of this data.) Furthermore, our purpose has been to understand the nature of traditional dual system estimation, and the consequences of alternate synthetic assumptions. For the most part the nature of the April 2001 A.C.E. data in relation to the census is analogous to that between earlier censuses and their dual system surveys. (In particular, both the 2000 census counts and the 2001 A.C.E. data contain correspondingly significant numbers of duplicates, such as presumably existed in earlier census data even though there was no way to explicitly identify them. See Section 2 on imputation for discussion

of one difference between 2000 and earlier censuses.) Furthermore the analysis of A.C.E.II involves a number of special complications and assumptions beyond those of the standard dual system analyses.

2. Imputation

We use *II*, the Census Bureau’s notation, to denote the number of imputations. Technically *II* is referred to as “insufficient information”. It is not unusual for some census records to contain incomplete information to a modest extent. If all or nearly all relevant information is missing so that the matching of the P-sample records to the E-sample enumerations is not feasible, then the record is described as having insufficient information. Here we use the word “imputation” generally to describe records that for some reason do not include enough information to be included in the A.C.E. process. Broadly speaking, census imputation also includes imputation for item non-response for records in the A.C.E., and imputation for matching status in the A.C.E. process. Yet in our context, imputation is referred to as the whole records not included in the A.C.E. process due to insufficient information. In the 2000 census, imputations included two parts: inherent imputation and late adds.

One can identify two basic kinds of inherent imputation. Sometimes we do know with reasonable certainty how many people there are in the household, but lack personal information about them as is needed for the matching of the E-sample and the P-sample in the dual system process. In this case, we just need to impute demographic information for each person. On the other hand, sometimes the actual number of people in the household is also unknown. In this circumstance both the true counts and personal information need to be imputed. It is even possible to give a finer subdivision of types of inherent imputations. See Norwood and Citro [11].

Imputation related to a large number of late-adds was a special feature of the 2000 census. Because of its concern about address duplication, the Census Bureau created a special research program just after the basic census data was collected. The Bureau was able to identify, and pulled out, approximately 6 million person records in 2.4 million housing units as potential duplicates. Later on, approximately 2.4 million persons in 1 million housing units were reinstated into the census. However, this was too late for the 2.4 million people to be included in the A.C.E. process. Hence they were referred to as “Late Adds” and were treated similarly to imputation data. For details of research on duplicates, see ESCAP [4]. Table 1 is a comparison of the distributions of imputation in 1990 and 2000.

Besides the fact that there was no special treatment for Late Adds in the 1990 census, there is a significant difference in terms of the ratio of imputations from households with known person count and imputations from households with unknown person count between the 1990 and the 2000 Census. In 2000, that ratio was about 4. Yet in 1990, the ratio was 44 which is 10 times larger than that in 2000.

TABLE 1
Number of imputations (II) as a percentage of census count (C)

Imputation type	2000 Census	1990 Census
Known Person Count	1.68	0.88
Unknown Person Count	0.43	0.02
Late Adds	0.85	0.00
Total	2.96	0.90

(Source: *The 2000 Census: Interim Assessment*)

The percentages of II from the 1980 census were more similar to those of 2000 than were the 1990 percentages.

In this paper, the item $C-II$ denotes the number of people with full information. They are frequently referred to as “data-defined” persons, and we use DD to denote the number of them in the following sections.

3. Dual system estimation

As we introduced before, the 2000 A.C.E. data consists of the E-sample and the P-sample. Based on the information of the E-sample and the P-sample, a dual system estimate of the population is produced for special subgroups, called post-strata. These post-strata estimates are then apportioned and recombined so as to form estimates for any geographic area, such as state, county, census block etc. We now discuss some aspects of this procedure.

3.1. Post-stratification

For the purpose of analysis, the population is divided into certain groups called post-strata. Sixty-four post-stratum groups were created based on information about geographic location, race, Hispanic origin, housing tenure etc. In addition there were 7 age/sex categories. Thus originally there were 448 post-strata. Later on, some small post-strata were collapsed together to form 416 final post-strata. [See Table 5 in the Appendix for details of the construction of post-strata.]

3.2. Dual system estimation

The dual system estimate for post-stratum i can be written as

$$(1) \quad D\widehat{SE}_i = DD_i \times \widehat{CR}_i \times \frac{1}{\widehat{MR}_i}.$$

Here DD_i is the number of data-defined persons in post-stratum i . \widehat{CR}_i and \widehat{MR}_i are the estimates of the E-sample correct enumeration rate and the P-sample matching rate respectively.

In the E-sample, enumerations are divided into two categories: correct enumerations and erroneous enumerations. The correct enumeration rate measures the accuracy of the census. It is estimated as

$$(2) \quad \widehat{CR}_i = \frac{CE_i}{CE_i + EE_i},$$

where CE_i denotes the number of correct enumerations and EE_i denotes the number of erroneous enumerations in post-stratum i .

The P-sample persons are taken into a matching procedure to see whether they can be matched with persons in the E-sample. The P-sample matching rate then measures the coverage of the census. The formula for \widehat{MR}_i is more complicated than that for the other elements of (1), and it is not particularly pertinent to the current considerations. The reader should consult Hogan [8] for details.

Since it was adopted by the Census Bureau to estimate the population, the dual system estimation method has been considered in principle a large-scale capture-recapture procedure. It can be motivated from an over-simplified, primitive model

for capture-recapture estimation. In this model, the interrelation of the P-sample and the E-sample can be schematically summarized in a two by two table, and elements in the two by two table are estimated based on the assumption of the independence of the E-sample and the P-sample. For a detailed overview of dual system estimation, see Hogan [7].

3.3. Synthetic assumption

The census provides population figures for geographic subdivisions much smaller than those defined by post-stratum boundaries. These “smaller areas” include states, congressional districts, metropolitan areas, and even divisions as small as census tracts and census blocks within tracts.

In order to get smaller area estimates, the estimates \widehat{DSE}_i for each post-stratum must be divided up and apportioned to geographic areas lying within that post-stratum. This procedure is called synthetic estimation and the assumption(s) that support its validity is (are) referred to as the synthetic assumption.

It seems to us that there are various reasonable forms of synthetic assumptions that could be proposed, and these lead in practice to different smaller area population estimates. For now we first present the formula implemented by the Bureau. Then we later contrast it with alternative formulas that also seem to us to be plausible.

For the purpose of synthetic estimation, the Census Bureau assumes that the estimate, \widehat{DSE}_i , should be divided in proportion to the total census counts within its post-stratum. Let the index k , $k = 1, 2, \dots, K_i$ refer to geographic subregions within post-stratum i . Let C_{ik} denote the total census counts for post-stratum i and region k , and let C_i denote the totals for the post-stratum. The Bureau population estimate for post-stratum i region k is then called \widehat{DSE}_{ik} or S_{ik} and is given by the formula

$$(3) \quad S_{ik} \equiv \widehat{DSE}_{ik} = \frac{C_{ik}}{C_i} \widehat{DSE}_i.$$

This reflects the Bureau’s synthetic assumption that the population distribution for smaller areas within a post-stratum is homogeneous with respect to the census counts for those areas within that post-stratum.

Formula (3) is often rephrased in a different but equivalent format. Define the Coverage Correction Factor for post-stratum i (CCF_i) by

$$(4) \quad CCF_i = \frac{\widehat{DSE}_i}{C_i}.$$

Then

$$(5) \quad S_{ik} = C_{ik} CCF_i.$$

There is a different but equivalent way to interpret (3) or (5). The Census Bureau’s estimate can also be written as

$$(6) \quad S_{ik} = C_{ik} + (\widehat{DSE}_i - C_i) \times \frac{C_{ik}}{C_i}.$$

We will later build upon this interpretation.

In summary, for geographic region k this gives the following population estimate:

$$(7) \quad S_k = \sum_i S_{ik} = \sum_i C_{ik} CCF_i.$$

Here in (7), S_k is called the synthetic dual system estimate, abbreviated as SynDSE. It is clear from its definition that it applies the same adjustment factor for people in each post-stratum, and aggregates the adjusted post-stratum level population numbers for an estimate of the population of the entire geographic area.

3.4. Rationale for post-stratification

The preceding discussion highlights one main rationale and target for post-stratification. Accuracy of the synthetic estimation formula (3) rests on the assumption that the population for the geographic areas within post-strata is distributed in proportion to the census count.

There are at least two other reasons for post-stratification in connection with dual system estimation. The logic supporting the dual system estimate requires that the matching rate be constant for individuals within post-strata. Violation of this will, in general, lead to bias in the dual system estimate (1) of the post-stratum population. Such a situation is referred to as “correlation bias”. There are many discussions of correlation bias in the literature. For example, Seker and Deming [12] had an early discussion on correlation bias. Bell [1] introduced a third system to estimate the correlation bias. Freedman and Wachter [5] also had a discussion on correlation bias and heterogeneity. Zhao [14] investigated the data of the 2000 census to test the plausibility of the assumption of absence of correlation bias.

A third, though perhaps less important, rationale for post-stratification is that, in principle, suitably chosen post-strata can reduce the variance of estimates given through formulas such as (1) and (3). Conversely, a choice of too many post-strata with consequently small sample sizes within each post-stratum can lead to estimators with inflated variances. See Hogan [7] for a discussion of this in relation to the 1990 census. See Freedman and Wachter [6] for a perspective on post-stratification and its effects in the 2000 census.

4. Alternative formulas

In this section, we present three alternative formulas for synthetic estimation. The Census Bureau’s formula is based on the synthetic assumption that the population distribution for small areas within a post-stratum is homogeneous with respect to the census counts (including imputations) for those areas within that post-stratum. Our alternative formulas are sensitive to the the homogeneity of imputations in the census, and its role in the synthetic estimation of subpopulation counts.

4.1. First alternative formula

Note that the estimates \widehat{DSE}_i are computed only from enumerations of data-defined people. That is because C_i does not appear in (1). Thus the estimates of \widehat{DSE}_i of post-stratum totals involve DD directly, but do not involve the number of counts labelled as II . It can thus be plausibly argued that the counts II should also not play a role in distributing \widehat{DSE}_i geographically within post-strata.

As noted in Section 3.4, homogeneity assumptions relative to the components of (1) are already part of the general justification for dual system estimation. From this perspective, it also seems reasonable to assume that the population for the geographic area within post-strata should be proportional to the enumeration of data defined people. This form of synthetic assumption leads to the alternate estimate S_{ik}^1 described as the formula

$$(8) \quad S_{ik}^1 \equiv D\widehat{SE}_{ik}^1 = \frac{DD_{ik}}{DD_i} D\widehat{SE}_i,$$

where DD_{ik} is the number of data-defined persons in geographic region k within post-stratum i , $i = 1, 2, \dots, I$, $k = 1, 2, \dots, K_i$.

There is another way to view the formula for S_{ik}^1 . For each post-stratum i , consider DCF_i (Data-defined Coverage Factor) as a replacement of CCF_i . Their relationship is described in the following formula

$$(9) \quad DCF_i = \frac{C_i}{C_i - II_i} = \frac{C_i}{DD_i} CCF_i.$$

Then applying the same Data-defined Coverage Factor for post-stratum i to the number of data-defined persons in geographic region k within post-stratum i , the corresponding S_{ik}^1 for geographic level k is thus written as

$$(10) \quad S_{ik}^1 = DD_{ik} DCF_i.$$

Note that (9) implies that $DCF_i = \frac{D\widehat{SE}_i}{DD_i}$, it is easy to show that (8) and (10) are equivalent.

4.2. Second alternative formula

It can be plausibly argued that the distribution of imputations $II_{ik} = C_{ik} - DD_{ik}$, $k = 1, 2, \dots, K_i$ is a valid reflection of distribution of the true undercount relative to C_{ik} within the post-stratum. Presumably imputations are concentrated in areas where it is intrinsically hard to count people, and hence areas with high undercount rate would be expected to have high imputation rate. Since the “true” undercount is not observed, it is hard, or impossible to devise a way to check this assertion. If it were valid, then the desirable estimate for the true population would be derived by distributing the post-stratum undercount estimates within the post-stratum in proportion to II_{ik} . This leads to the formula

$$(11) \quad \begin{aligned} S_{ik}^2 &= C_{ik} + (D\widehat{SE}_i - C_i) \times \frac{II_{ik}}{II_i} \\ &= C_{ik} + (DD_i \times DCF_i - C_i) \times \frac{II_{ik}}{II_i}. \end{aligned}$$

As we noted before, the estimate of the total undercount for post-stratum i is $D\widehat{SE}_i - C_i$, and this undercount is distributed to each geographic level proportionally to its imputation rate within the post-stratum. The estimate for the population is then the census counts plus the estimated undercount. In summary, this formula is the same as (6) except that $\frac{II_{ik}}{II_i}$ is substituted for $\frac{C_{ik}}{C_i}$.

4.3. Third alternative formula

Note that the Census Bureau's formula (6) is $S_k = C_{ik} + (\widehat{DSE}_i - C_i) \times \frac{C_{ik}}{C_i}$. Compare this with (11), and another reasonable formula comes out naturally as

$$\begin{aligned} (12) \quad S_{ik}^3 &= C_{ik} + (\widehat{DSE}_i - C_i) \times \frac{DD_{ik}}{DD_i} \\ &= C_{ik} + (DD_i \times DCF_i - C_i) \times \frac{DD_{ik}}{DD_i}. \end{aligned}$$

In words, this formula begins from a base of the census counts C_{ik} (including II_{ik}). It then considers the distribution of DD_{ik} as a reflection of the true undercount rate at geographic level within post-strata.

Clearly all of the formulas presented here have the same normalization property

$$(13) \quad \sum_k S_{ik}^l = \sum_k S_{ik} = \widehat{DSE}_i, \quad l = 1, 2, 3.$$

Also, if we take the summation over post-stratum index i , then we will have the estimate of the population at geographic area k as

$$(14) \quad S_k^l = \sum_i S_{ik}^l, \quad l = 1, 2, 3.$$

5. Results from alternative formulas at state level

5.1. Comparison of shares at state level

Allocating seats in the House of Representatives is the original constitutional mandate for which the decennial census was established. Much attention was put on which states had gained or lost seats. It is of primary interest to compare different formulas at the state level.

Figure 1 shows comparison of alternative formulas and the Census Bureau's formula for the 16 largest states. [See Figure 5 in the Appendix for the full comparison of all 51 states.] The comparison is made in the sense of population shares. A state's population share is normally defined as its percentage of the national total. Thus they do not affect estimates for national totals. The horizontal line for each state shows the confidence interval of share difference: SynDSE (S_k) share minus census share. The standard error of share difference is computed from Davis [3] published by the Census Bureau. The square represents the share difference between S_k and census, the dot represents the share difference between S_k^1 and census, and the triangle represents the share difference between S_k^2 and census. The share difference between S_k^3 and census is omitted from the figure since it is very close to the one between S_k and census.

The most prominent feature is for the state of New York where the difference calculated from S_k^1 falls very far outside of (below) the confidence interval calculated from census formula. For several other states the result for S_k^1 is also outside the confidence interval (above, as for North Carolina, Virginia, and Ohio, or below, as for Indiana and Illinois). S_k^2 agrees better with the census formula. For several large states, such as Texas, California, Florida and Pennsylvania, the square and the triangle are very close to each other. The result for New York is driven towards 0, although it still falls outside (above) the confidence interval.

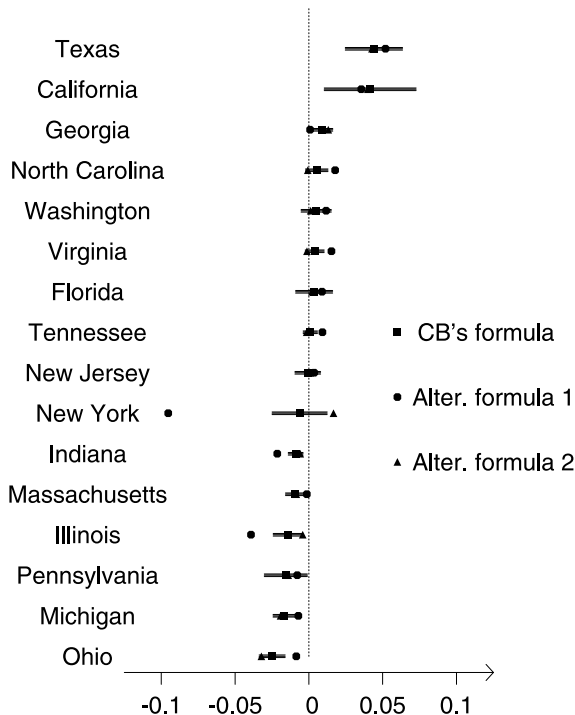


FIG 1. State level shares comparison from different formulas.

Interestingly, most of the time, the share difference of S_k and census falls between the difference of S_k^1 and census, and the difference of S_k^2 and census. This tells us that, in a sense the census formula is a compromise of the two alternatives we introduced.

5.2. Role of imputation

Imputations create the primary difference in practice between the Bureau's synthetic formula (3) and alternative formulas such as our (8), (11) and (12). Note that the assumption justifying (8) is that the undercount is homogeneous with respect to DD_{ik} for regions within post-strata. In contrast, the assumption justifying (3) is that of homogeneity with respect to $C_{ik} = DD_{ik} + II_{ik}$. If the imputation rates were stochastically homogeneous with respect to C_{ik} , then both formulas would have the same expectation, and would generally yield very similar results in practice.

Imputation rates for the 16 large states of Figure 1, together with the population shares from the census, are given in Table 2. In this table, the total imputation rates, the imputation rates from late adds (LA) and non late adds (Non-LA), as well as the census shares are listed. [See Table 6 in the Appendix for the full table for all 51 states.] The overall imputation rate for New York is considerably larger than the national rate of 3%.

Furthermore, what really matters is the imputation rates within post-strata within the state relative to those post-strata results elsewhere. Because of this it seems informative to supplement the overall imputation rates given in the table

TABLE 2
Imputation rates for the 16 states

State	II(Tot)	II(Non-LA)	II(LA)	Census Share	Number of post-strata	Mean II(Tot) of post-strata
NY	4.913	3.201	1.712	6.724	256	5.511
TX	3.508	2.633	0.875	7.417	264	3.962
IL	3.383	2.469	0.914	4.422	246	4.380
GA	3.300	2.349	0.951	2.907	234	3.994
CA	3.255	2.720	0.535	12.08	260	4.360
NJ	2.869	2.008	0.861	3.004	194	4.849
NC	2.795	1.640	1.156	2.849	236	3.558
IN	2.700	2.202	0.498	2.157	246	4.106
FL	2.672	2.113	0.558	5.700	236	4.534
MA	2.468	1.558	0.909	2.240	253	3.239
TN	2.465	1.599	0.867	2.025	222	3.372
WA	2.407	1.894	0.513	2.105	236	3.410
PA	2.322	1.574	0.748	4.331	242	3.595
VA	2.283	1.555	0.727	2.503	260	3.536
MI	1.876	1.341	0.536	3.541	260	3.039
OH	1.680	1.123	0.557	4.040	222	2.699

with per post-strata averages. As a result, Table 2 also gives the mean imputation rate per post-strata within state as computed from the following formula:

$$(15) \quad MIR_k = \frac{1}{n_k^*} \sum_{\{i, C_{ik} \neq 0\}} \frac{II_{ik}}{C_{ik}} \times 100\%$$

where n_k^* is the number of post-strata within the state with non-zero census counts, which is also listed in the table. Even a cursory examination of these imputation rates in the census reveals that an assumption for the imputations of stochastic homogeneity within post-strata is not reasonable. (A valid, formal test of this statistical hypothesis can be derived using the methods of Zhao [14]. This test decisively rejects the null hypothesis of stochastic homogeneity, with a p -value < 0.0001 .)

In Table 2, the comparison of New York and New Jersey points to an interesting phenomenon. Overall New Jersey has an imputation rate of 2.869%. This is fairly close to the national average. But it shares a lot of post-strata with New York. The mean value of the imputation rates per post-strata in New Jersey is 4.849%. This is the second highest among the 16 states. Yet as shown in Figure 1, in contrast to New York, the differences for New Jersey using S_k^1 and S_k^2 are quite close to that using the Census Bureau's S_k . The result is that although New Jersey has relatively high mean imputation rate per post-strata, its population estimate is not increased as much by the dual system as this might seem to warrant. One explanation for this is that an important neighboring state (New York) has even higher imputation rates.

From another point of view, we can consider our alternative formula one as a basic rate for estimate of population, while the Census Bureau's formula can be viewed as an attempt to use imputations with the hope of improving these basic estimates.

6. Results from alternative formulas at county-group level

To better investigate the differences among all the formulas, we conduct a further analysis down to a finer level: county-group level. Ideally our analysis might have been performed on the level of congressional districts. However we had only county

level data to work with. Hence we created county groups to roughly approximate the size and geographic contiguity of congressional districts. (In some cases our county groups were much more populous than congressional districts since we could not split counties into smaller districts.) In general, small adjacent counties are lumped to form a group with population roughly like a congressional district, while relatively large counties (for example, a county contains several congressional districts) would make a county-group by themselves. Totally we created 369 county-groups, on average each having 730,000 people.

For each county-group, an adjusted estimate (*SynDSE*) is constructed by the Census Bureau's formula and our alternative formula 1, 2 and 3. It seems most suitable to compare the adjustments to the relative shares. This is consistent with the discussion in Brown et al. [2] and Freedman and Wachter [5]. However we found direct statements of share differences to be less suitable in part because of unfamiliarity with the county-groups and variability in their sizes. Hence it seems more informative to express the adjustments in percentage terms from a base of the original census numbers. It can be easily shown that this measure is a linear transformation of the share difference, and as noted in the above references, the results from the percent adjustment would be consistently comparable to the share difference.

There are two possible choices of the base of the original census numbers. Naturally people would consider the census counts, and the relative percent difference can be expressed as

$$(16) \quad \text{reldif}^c = \frac{\text{SynDSE} - C}{C} \times 100\%.$$

However, one of the implications of the alternative formula one is that the number of data-defined person *DD* is a more basic quantity. Therefore we use *DD* as the base, and the relative percent difference is then defined as

$$(17) \quad \text{reldif}^d = \frac{\text{SynDSE} - DD}{DD} \times 100\%.$$

To account for the implication of imputation, (17) can be modified to be a measure called state adjusted difference (*SAD*), which is defined by

$$(18) \quad \text{SAD}_j = \left(\frac{\text{SynDSE}_j - DD_j}{DD_j} - \frac{II_s}{DD_s} \right) \times 100\%.$$

In (18), *j* is the county-group index, *s* is the state index. The following Table 3 illustrates the descriptive statistics for *SAD* using different formulas.

As we already found from the last section, the alternative formula three gives very similar results as the Census Bureau's. It is also noticeable from the table that overall there is no substantial difference in terms of the mean value of differences. [The results from *reldif*^c can be found in Table 7 in the Appendix, and they will give similar relative conclusions among county groups within a state.]

TABLE 3
Distribution of state adjusted difference at county group level

	Min	Max	Median	Mean	SD
CB's formula	-2.97	7.38	0.98	1.14	1.40
Alter. formula 1	-2.95	4.93	1.20	1.18	1.10
Alter. formula 2	-3.08	9.80	0.81	1.15	1.60
Alter. formula 3	-2.98	7.29	0.99	1.14	1.39

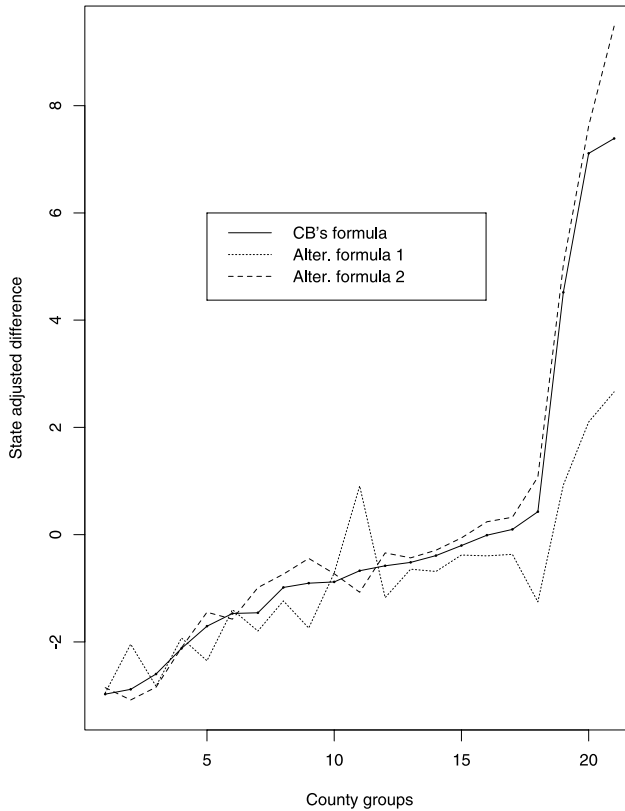


FIG 2. *State adjusted difference – New York (DD base).*

It is impossible to visually show the results of *SAD* from all county-groups in one figure; instead we illustrate the results in the following three states:

1. New York: because of the large discrepancy in share comparison (Figure 1) and the relatively large size (3rd biggest state)
2. New Jersey: because of the interesting phenomenon discussed in Section 5.2
3. California: because of the relatively large size (biggest state)

Figure 2 is the plot of *SAD* in each county-group in New York. [The table generating this figure can be found in the Appendix.] Each one of the 21 points on the X-axis represents a county group, and the state adjusted differences represented on the Y-axis are connected by a line. Different types of lines represent different formulas. Again, the results from alternative formula 3 are not shown in the figure because they are very close to those from the Census Bureau's formula. It is obvious that for the three counties in New York city (Bronx, Kings and Queens) which have a very large percent of imputation, the differences are much higher than those from other county-groups.

Figure 3 is the plot of *SAD* in each county-group in New Jersey. Despite the fact that New Jersey shares a lot of post-strata with New York, the scale of the differences is much smaller than that from New York.

Figure 4 is the plot of *SAD* in each county-group in California. From all three figures, it can be seen that most of the time, the lines using Census Bureau's formula lie between the lines using our alternative formula 1 and alternative formula 2.

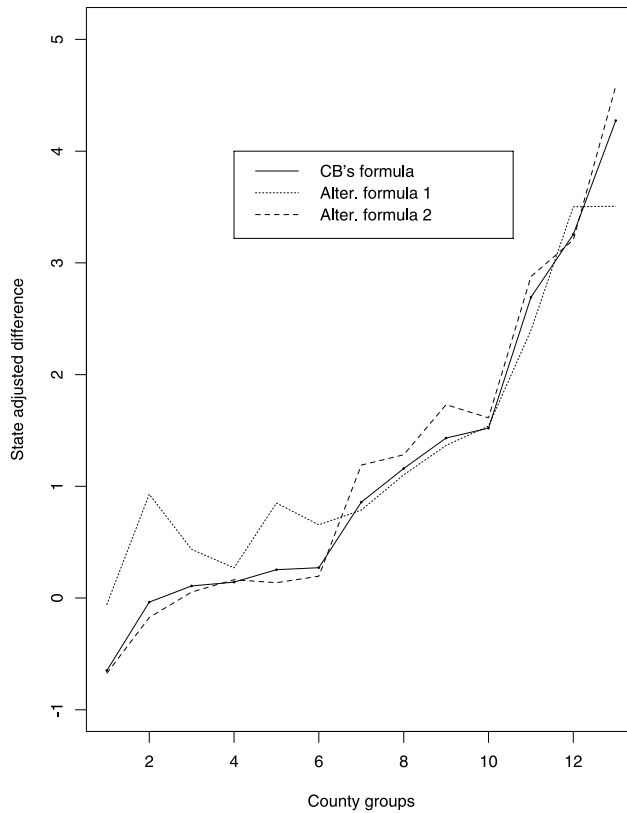


FIG 3. State adjusted difference – New Jersey (*DD* base).

This confirms that the Census Bureau's formula is kind of a compromise of the two alternatives.

It can also be seen that in general, at the lower end of the figure (smaller difference between *SynDSE* and *DD*), the difference using Census Bureau's formula tends to be lower (higher) than that using alternative formula 1 (using alternative formula 2), while at the upper end of the figure (larger difference between *SynDSE* and *DD*), the difference using Census Bureau's formula tends to be higher (lower) than that using alternative formula 1 (using alternative formula 2). (The detailed results at each county group in these three states could be found in Table 8 through Table 10 in the Appendix.)

7. Comparison of different formulas

7.1. Comparison of four formulas

As stated earlier, if the imputation rates were stochastically homogeneous with respect to the census count, then all the formulas would have the same expectation.

It is easy to prove that if $\frac{II_{ik}}{II_i} = \frac{C_{ik}}{C_i}$, then $S_k^1 = S_k^2 = S_k^3 = S_k$.

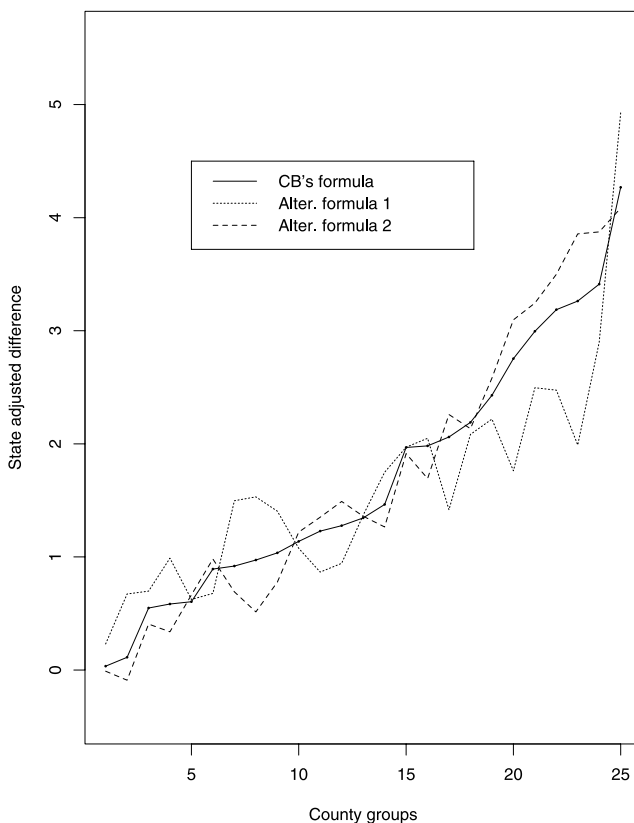


FIG 4. State adjusted difference - California (DD base).

7.2. When is DCF better

Our alternative formula (10) uses *DCF* instead of *CCF*. One may wonder under which conditions does *DCF* behave better than *CCF*.

Consider the following simpler case: there are two states for a single post-stratum, and there are no people who moved between the census day and the A.C.E. interview. The corresponding counts in state 1 and 2 within post-stratum are: $CE_1, CE_2, EE_1, EE_2, MN_1, MN_2, NN_1, NN_2, II_1, II_2$, and they are all observable. Here CE_j, EE_j, MN_j, NN_j , and II_j ($j = 1, 2$) denotes the number of correct enumerations, erroneous enumerations, matched non-movers, unmatched non-movers, and imputations respectively. For a formal definition of these types of counts, see Norwood and Citro [11]. As also shown in Norwood and Citro [11], *CCF* and *DCF* can be written as functions of these five types of counts

$$(19) \quad CCF = \frac{CE_1 + CE_2}{CE_1 + CE_2 + EE_1 + EE_2 + II_1 + II_2} \times \frac{NN_1 + NN_2}{MN_1 + MN_2},$$

$$(20) \quad DCF = \frac{CE_1 + CE_2}{CE_1 + CE_2 + EE_1 + EE_2} \times \frac{NN_1 + NN_2}{MN_1 + MN_2}.$$

To further simplify the case, we assume that the two states are equal in size, i.e. $CE_1 = CE_2, MN_1 = MN_2, NN_1 = NN_2$

The following analysis makes a comparison of the squared errors resulting from use of (3) and (10). In order to make this comparison it is necessary to make some assumptions about the true population. The analysis is somewhat simple under the plausible assumption that the unbiased DSE from the two by two tables within each state describes the true population parameters. A similar analysis is possible under other assumptions.

The unbiased DSE from the actual two by two tables within each state can be written as

$$S_1^t = CE_1 \times \frac{NN_1}{MN_1} = S, \quad S_2^t = CE_2 \times \frac{NN_2}{MN_2} = S.$$

The synthetic DSEs for state 1 and 2 within post-stratum calculated from *CCF* and *DCF* (use alternative formula one) are

$$(21) \quad S_i^c = \frac{CE_i + EE_i + II_i}{CE_1 + CE_2 + EE_1 + EE_2 + II_1 + II_2} \times 2S, \quad i = 1, 2,$$

and

$$(22) \quad S_i^d = \frac{CE_i + EE_i}{CE_1 + CE_2 + EE_1 + EE_2} \times 2S, \quad i = 1, 2.$$

Define the variance, i.e. the squared error of synthetic DSE from the true population, as

$$\begin{aligned} \Delta_c &= (S_1^c - S_1^t)^2 + (S_2^c - S_2^t)^2 \\ &= 2S^2 \left(\frac{CE_1 + EE_1 + II_1 - (CE_2 + EE_2 + II_2)}{CE_1 + EE_1 + II_1 + CE_2 + EE_2 + II_2} \right)^2, \\ \Delta_d &= (S_1^d - S_1^t)^2 + (S_2^d - S_2^t)^2 \\ &= 2S^2 \left(\frac{CE_1 + EE_1 - (CE_2 + EE_2)}{CE_1 + EE_1 + CE_2 + EE_2} \right)^2. \end{aligned}$$

The difference of Δ_d and Δ_c is

$$\begin{aligned} (23) \quad \Delta_d - \Delta_c &= 2S^2 \left\{ \left(\frac{EE_1 - EE_2}{2CE_1 + EE_1 + EE_2} \right)^2 \right. \\ &\quad \left. - \left(\frac{EE_1 + II_1 - (EE_2 + II_2)}{2CE_1 + EE_1 + EE_2 + II_1 + II_2} \right)^2 \right\} \\ &= 2S^2 \left\{ \left(\frac{EE_1 - EE_2}{2CE_1 + EE_1 + EE_2} + \frac{EE_1 + II_1 - EE_2 - II_2}{2CE_1 + EE_1 + EE_2 + II_1 + II_2} \right) \right. \\ &\quad \left. \left(\frac{EE_1 - EE_2}{2CE_1 + EE_1 + EE_2} - \frac{EE_1 + II_1 - EE_2 - II_2}{2CE_1 + EE_1 + EE_2 + II_1 + II_2} \right) \right\}. \end{aligned}$$

If $CE \gg (EE, II)$, as is usually the case, then

$$(24) \quad \Delta_d - \Delta_c \approx \frac{-2S^2(4CE_1(EE_1 - EE_2) + 2CE_1(II_1 - II_2))(2CE_1(II_1 - II_2))}{(2CE_1 + EE_1 + EE_2)^2(2CE_1 + EE_1 + EE_2 + II_1 + II_2)^2}.$$

TABLE 4
Frequency table of better performance of *DCF* among large/small post-strata

	CCF	DCF	Total
Small	48	89	137
Large	23	84	107
Total	71	173	244

From (24) we have

- If $EE_1 = EE_2$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
- If $II_1 = II_2$ then $\Delta_d - \Delta_c \geq 0$, *CCF* is better.
- If $EE_1 \neq EE_2$ and $II_1 \neq II_2$.
 - If $EE_1 > EE_2$ and $II_1 > II_2$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
 - If $EE_1 > EE_2$ and $II_1 < II_2$:
 - * If $EE_1 - EE_2 \leq -\frac{II_1 - II_2}{2}$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
 - * If $EE_1 - EE_2 > -\frac{II_1 - II_2}{2}$ then $\Delta_d - \Delta_c > 0$, *CCF* is better.

More generally, we assume $CE_2 = \lambda CE_1$, $MN_2 = \lambda MN_1$, $NN_2 = \lambda NN_1$, since homogeneity assumption appears to hold for the two largest groups: CE and MN. For the setup and results from the test of homogeneity assumption, see Zhao [14]. Similarly we have

- If $\lambda EE_1 = EE_2$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
- If $\lambda II_1 = II_2$ then $\Delta_d - \Delta_c \geq 0$, *CCF* is better.
- If $\lambda EE_1 \neq EE_2$ and $\lambda II_1 \neq II_2$.
 - If $\lambda EE_1 > EE_2$ and $\lambda II_1 > II_2$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
 - If $\lambda EE_1 > EE_2$ and $\lambda II_1 < II_2$:
 - * If $\lambda EE_1 - EE_2 \leq -\frac{\lambda II_1 - II_2}{2}$ then $\Delta_d - \Delta_c \leq 0$, *DCF* is better.
 - * If $\lambda EE_1 - EE_2 > -\frac{\lambda II_1 - II_2}{2}$ then $\Delta_d - \Delta_c > 0$, *CCF* is better.

The above discussion gives certain conditions when the Census Bureau's correction factor (4) or the alternative correction factor (9) performs better than the other one. To show the empirical results from the data, let's consider a simple case. Suppose we regard New York state as state 1, and all the other states together as state 2, then we calculate the *DCF* and *CCF* for the 244 post-strata that are in both states. We found that *DCF* is better in 70% of post-strata which exist in both state 1 and state 2. Furthermore, if we categorize the post-strata into two groups: large post-strata (having more than 50,000 correct enumerations) and small post-strata, *DCF* performs much better in the large post-strata.

From Table 4, it could be seen that *DCF* (corresponding to formula (10)) performs better about 65% of the time in small post-strata and 80% of time in large post-strata.

8. Conclusion

The major purpose of this paper is to better understand the 2000 A.C.E. process by providing alternative formulas. To construct these three formulas, alternate forms of the synthetic assumption are used, and the structure of imputation is analyzed. We find that the alternative estimation formulas seem also justifiable.

It is perhaps hard to tell which formula gives generally more accurate results. It appears to us that each one has its own merit and no one dominates another. In addition, there seems no way with existing data to compare the biases of the formulas. Nonetheless, it appears that the first of the alternatives would achieve smaller variance than that of the Census Bureau's formula if the number of erroneous enumerations and the number of imputations are positively correlated, which holds true in most of the cases.

What we do observe is that the Census Bureau's formula tends to be a compromise among the three alternatives. For this reason it seems to us reasonable to stick to the original one, especially in view of a lack of further evidence.

All the Census Bureau's formula and our alternative formulas use the total number of imputations to create population estimates. As noted in Section 2, there are different classes of imputation. It may be preferable to use only some subsets of imputations, and create formulas in different ways.

Finally we want to point out that the correct enumeration rate $CE/(CE + EE)$ is estimated in producing synthetic estimation. This estimate is another potential source of heterogeneity, and the related synthetic assumption on it should be studied. A valid, formal test of the hypothesis that the correct enumeration rate is geographically homogeneous within post-strata for states or counties can be derived using the methods of Zhao [14]. This test shows there is significant non-homogeneity. (The details of this test will be reported elsewhere.) It would be desirable to also see how this inhomogeneity affects synthetic estimates results. However, unlike II , the components CE and EE are not measured for the entire census, but rather only for the A.C.E. sample blocks. Thus it is unclear how to use existing data to create estimates related to this factor.

Appendix

TABLE 5. Schematic for post-stratification variables (see Section 3.1 for further description)
(MSA: Metropolitan Statistical Area; TEA: Type of Enumeration Area; MO/MB: Mail out/Mail back)

Race/Hispanic Origin	Tenure	MSA/TEA	High return rate				Low return rate			
			NE	MW	S	W	NE	MW	S	W
Domain 7:	Owner	Large MSA MO/MB	1	2	3	4	5	6	7	8
Non-Hispanic White		Medium MSA MO/MB	9	10	11	12	13	14	15	16
and Other		Small MSA & Non-MSA MO/MB	17	18	19	20	21	22	23	24
		All Other TEAs	25	26	27	28	29	30	31	32
	Non-Owner	Large MSA MO/MB		33				34		
		Medium MSA MO/MB		35				36		
		Small MSA & Non-MSA MO/MB		37				38		
		All Other TEAs		39				40		
Domain 4:	Owner	Large MSA MO/MB		41				42		
Non-Hispanic Black		Medium MSA MO/MB								
		Small MSA & Non-MSA MO/MB		43				44		
		All Other TEAs								
	Non-Owner	Large MSA MO/MB		45				46@		
		Medium MSA MO/MB								
		Small MSA & Non-MSA MO/MB		47				48		
		All Other TEAs								
Domain 5: Native Hawaiian or Pacific Islander		Owner					49			
		Non-Owner					50			
Domain 6:		Owner					51			
Non-Hispanic Asian		Non-Owner					52			
Domain 3:	Owner	Large MSA MO/MB		53					54	
Hispanic		Medium MSA MO/MB								
		Small MSA & Non-MSA MO/MB		55					56	
		All Other TEAs								
	Non-Owner	Large MSA MO/MB		57					58	
		Medium MSA MO/MB								
		Small MSA & Non-MSA MO/MB		59					60	
		All Other TEAs								
Domain 1: On Reservation American Indian or Alaska Native		Owner					61			
		Non-Owner					62			
Domain 6: Off Reservation American Indian or Alaska Native		Owner					63			
		Non-Owner					64			

TABLE 6
Imputation rates for 51 states

State	II(Tot)	II(Non-LA)	II(LA)	Census Share	Number of post-strata	Mean II(Tot) of post-strata
NY	4.913	3.201	1.712	6.724	256	5.511
NM	4.474	2.895	1.579	0.652	236	4.137
HI	4.247	2.913	1.334	0.430	222	4.781
WY	3.921	2.588	1.333	0.175	148	4.166
NV	3.918	3.257	0.661	0.718	236	4.479
AZ	3.891	3.145	0.746	1.835	236	4.942
VT	3.887	2.223	1.664	0.215	134	3.961
DC	3.860	3.726	0.134	0.196	144	4.373
TX	3.508	2.633	0.875	7.417	264	3.962
AL	3.491	2.212	1.279	1.584	235	4.104
IL	3.383	2.469	0.914	4.422	246	4.380
DE	3.339	2.901	0.438	0.277	222	4.900
RI	3.302	2.360	0.942	0.369	221	4.821
GA	3.300	2.349	0.951	2.907	234	3.994
CA	3.255	2.720	0.535	12.08	260	4.360
SC	3.221	2.145	1.076	1.417	236	4.423
MD	3.074	2.503	0.572	1.887	222	4.178
NH	3.056	1.987	1.070	0.439	217	4.440
MT	3.039	1.583	1.456	0.321	152	4.478
MS	3.038	1.677	1.360	1.005	228	3.289
LA	2.886	1.886	1.000	1.584	236	3.432
NJ	2.869	2.008	0.861	3.004	194	4.849
AR	2.810	1.403	1.407	0.950	222	2.852
NC	2.795	1.640	1.156	2.849	236	3.558
CO	2.786	2.039	0.747	1.535	236	4.119
IN	2.700	2.202	0.498	2.157	246	4.106
FL	2.672	2.113	0.558	5.700	236	4.534
ME	2.604	1.258	1.345	0.453	184	2.889
AK	2.584	1.385	1.199	0.202	152	3.131
ID	2.554	1.821	0.733	0.461	152	3.374
WV	2.506	0.856	1.651	0.645	215	2.625
MA	2.468	1.558	0.909	2.240	253	3.239
TN	2.465	1.599	0.867	2.025	222	3.372
KT	2.447	1.164	1.283	1.435	222	2.888
WA	2.407	1.894	0.513	2.105	236	3.410
CT	2.390	1.544	0.847	1.205	256	3.875
UT	2.369	1.765	0.604	0.801	236	2.972
SD	2.362	1.392	0.970	0.266	140	2.733
PA	2.322	1.574	0.748	4.331	242	3.595
VA	2.283	1.555	0.727	2.503	260	3.536
OK	2.261	1.282	0.979	1.220	236	2.517
OR	2.260	1.711	0.549	1.222	236	3.639
WI	2.153	1.600	0.553	1.903	258	4.311
MO	2.098	1.200	0.898	1.986	222	3.050
ND	1.985	1.003	0.983	0.226	138	2.342
KS	1.904	1.227	0.678	0.953	236	2.739
MI	1.876	1.341	0.536	3.541	260	3.039
MN	1.873	1.237	0.636	1.748	236	3.463
OH	1.680	1.123	0.557	4.040	222	2.699
IA	1.629	0.963	0.666	1.032	215	2.589
NE	1.608	0.994	0.615	0.607	236	2.487

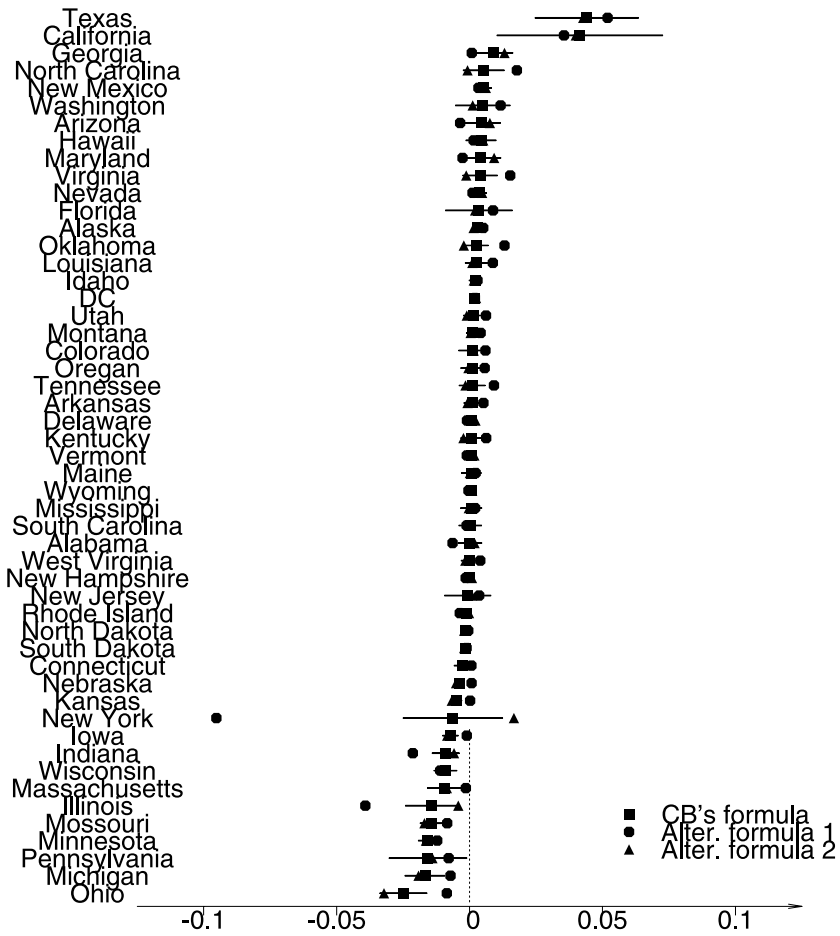


FIG 5. Share comparison at state level.

TABLE 7
Distribution of relative difference between census and SynDSE at county group level

	Min	Max	Median	Mean	SD
CB's formula	-0.13	2.96	1.15	1.14	0.54
Alter. formula 1	-3.26	3.78	1.22	1.18	0.84
Alter. formula 2	-0.14	4.31	1.06	1.13	0.68
Alter. formula 3	-0.13	2.97	1.15	1.14	0.54

TABLE 8
County group level results in New Jersey
(Through Table 8 to 10, the second column "CB's" lists the results using the Census Bureau's formula, the third column "Alter. 1" lists the results using alternative formula 1, and the fourth column "Alter. 2" lists the results using alternative formula 2.)

<i>Relative difference in New Jersey (census as the base)</i>					
Counties	CB's	Alter. 1	Alter. 2	Census	II/Census
Passaic	1.566	1.282	1.743	479073	3.863
Essex	1.471	1.706	1.423	770844	4.462
Hudson	1.470	0.745	1.763	599525	5.369
Somerset, Union	1.250	1.268	1.339	807714	3.087
Atlantic, Cape May & Cumberland, Salem	1.234	1.179	1.354	542964	2.766
Mercer	1.223	1.160	1.514	329669	3.030
Middlesex, Monmouth	1.123	1.707	1.009	1334607	2.021
Morris	1.064	1.387	1.011	461026	1.938
Sussex, Warren	0.973	1.920	0.838	243450	1.890
Bergen	0.967	1.344	0.894	872769	2.187
Burlington, Ocean	0.964	1.089	0.985	912247	2.068
Camden, Gloucester	0.951	0.882	1.274	747998	2.756
Hunterdon	0.797	1.380	0.772	117643	1.474

<i>State adjusted difference in New Jersey</i>					
Counties	CB's	Alter. 1	Alter. 2	DD	II/DD
Hudson	4.273	3.507	4.583	567337	5.674
Essex	3.255	3.502	3.206	736452	4.670
Passaic	2.693	2.398	2.877	460565	4.019
Somerset, Union	1.521	1.540	1.613	782780	3.185
Mercer	1.432	1.367	1.732	319680	3.125
Atlantic, Cape May & Cumberland, Salem	1.159	1.103	1.282	527948	2.844
Camden, Gloucester	0.858	0.787	1.190	727384	2.834
Bergen	0.271	0.656	0.196	853681	2.236
Middlesex, Monmouth	0.254	0.851	0.138	1307639	2.062
Burlington, Ocean	0.143	0.270	0.164	893380	2.112
Morris	0.108	0.437	0.054	452090	1.977
Sussex, Warren	-0.036	0.930	-0.174	238849	1.926
Hunterdon	-0.649	-0.057	-0.674	115909	1.496

TABLE 9
County group level results in New York

<i>Relative difference in New York (census as the base)</i>					
Counties	CB's	Alter. 1	Alter. 2	Census	II/Census
Bronx	2.405	-1.893	4.313	1285415	9.016
Clinton, Franklin, Fulton, Hamilton & Jefferson, Lewis Oswego, St Lawrence	1.683	1.512	1.819	524735	3.126
Chenango, Delaware, Herkimer & Madison, Oneida, Otsego, Schoharie	1.486	1.364	1.568	530826	3.023
Broome, Sullivan, Tioga, Tompkins, Ulster	1.480	1.193	1.576	562836	3.147
New York	1.477	3.012	1.088	1477358	2.887
Allegany, Cattaraugus, Chautauqua & Chemung, Schuyler, Steuben, Yates	1.280	1.036	1.522	484489	2.786
Dutchess, Putnam	1.263	0.891	1.506	355568	3.702
Kings	1.255	-3.263	1.717	2426027	9.817
Orange, Rockland	1.212	0.764	1.430	606779	3.851
Columbia, Essex, Greene, Rensselaer & Saratoga, Warren, Washington	1.095	0.520	1.326	603542	3.339
Westchester	0.999	1.162	1.149	899806	3.150
Albany, Montgomery, Schenectady	0.999	1.067	0.896	469399	2.602
Queens	0.945	-2.359	1.417	2202506	7.967
Cayuga, Cortland, Onondaga	0.794	0.986	0.807	567471	2.184
Nassau	0.736	-0.072	1.183	1312886	3.380
Monroe	0.735	1.566	0.541	708834	1.512
Erie	0.684	0.057	0.939	919474	2.683
Niagara, Orleans	0.532	0.311	0.291	256313	1.984
Suffolk	0.491	0.161	0.947	1390791	3.103
Genesee, Livingston, Ontario & Seneca, Wayne, Wyoming	0.252	0.273	0.372	376399	1.899
Richmond	-0.035	-1.626	0.576	434542	5.332
<i>State adjusted difference in New York</i>					
Counties	CB's	Alter. 1	Alter. 2	DD	II/DD
Bronx	7.386	2.662	9.482	1169523	9.909
Kings	7.110	2.100	7.622	2187875	10.885
Queens	4.517	0.927	5.030	2027022	8.657
Richmond	0.428	-1.252	1.073	411372	5.632
Orange, Rockland	0.098	-0.367	0.325	583412	4.005
Dutchess, Putnam	-0.011	-0.397	0.241	342405	3.844
Clinton, Franklin, Fulton, Hamilton & Jefferson, Lewis Oswego, St Lawrence	-0.203	-0.379	-0.062	508331	3.227
Broome, Sullivan, Tioga, Tompkins, Ulster	-0.390	-0.686	-0.291	545126	3.249
Chenango, Delaware, Herkimer & Madison, Oneida, Otsego, Schoharie	-0.517	-0.643	-0.433	514779	3.117
Columbia, Essex, Greene, Rensselaer & Saratoga, Warren, Washington	-0.580	-1.175	-0.340	583388	3.455
New York	-0.673	0.908	-1.073	1434701	2.973
Westchester	-0.882	-0.714	-0.728	871460	3.253
Nassau	-0.906	-1.742	-0.443	1268496	3.499
Allegany, Cattaraugus, Chautauqua & Chemung, Schuyler, Steuben, Yates	-0.985	-1.236	-0.736	470993	2.865
Suffolk	-1.457	-1.798	-0.987	1347631	3.203
Albany, Montgomery, Schenectady	-1.470	-1.400	-1.575	457183	2.672
Erie	-1.708	-2.351	-1.445	894808	2.757
Cayuga, Cortland, Onondaga	-2.123	-1.927	-2.109	555079	2.232
Niagara, Orleans	-2.601	-2.826	-2.846	251228	2.024
Monroe	-2.885	-2.042	-3.082	698113	1.536
Genesee, Livingston, Ontario & Seneca, Wayne, Wyoming	-2.974	-2.953	-2.852	369250	1.930

TABLE 10
County group level results in California

<i>Relative difference in California (census as the base)</i>					
Counties	CB's	Alter. 1	Alter. 2	Census	II/Census
Imperial	2.959	3.589	2.783	131317	4.343
Kings	2.384	1.891	2.827	109332	4.114
San Luis Obispo, Santa Barbara	2.191	2.257	1.911	613840	2.995
Monterey, San Benito, Santa Cruz	2.086	1.607	2.325	680087	4.018
Merced, Stanislaus	2.050	1.848	2.198	647207	3.538
Del Norte, Humboldt, Lake, Mendocino, Napa	1.918	1.920	1.865	406509	3.242
Kern, Tulare	1.913	1.233	2.213	993655	4.352
Los Angeles	1.829	1.727	1.776	9344086	3.529
Butte, Lassen, Modoc, Nevada, Plumas & Shasta, Sierra Siskiyou, Trinity, Yuba	1.801	2.345	1.353	621777	2.431
Fresno, Madera, Mariposa	1.661	0.448	2.228	912453	4.657
Colusa, Glenn, Sutter, Tehama, Yolo	1.577	1.935	1.325	338148	2.704
San Francisco	1.572	0.623	1.900	756976	4.283
Inyo, San Bernardino	1.542	1.817	1.349	1682190	3.135
Alameda	1.408	1.969	1.186	1416006	2.757
San Joaquin	1.390	0.771	1.581	544827	3.827
Riverside	1.380	1.404	1.395	1511034	3.179
Santa Clara	1.282	1.223	1.361	1652871	3.081
Orange	1.275	0.953	1.483	2803924	3.216
San Diego	1.228	1.623	0.989	2716820	2.616
San Mateo	1.192	0.841	1.311	696711	3.252
Ventura	1.131	1.150	1.189	739985	2.729
Sacramento	1.105	1.250	0.966	1198004	2.702
Contra Costa, Solano	1.065	1.612	0.868	1316047	2.330
Alpine, Amador, Calaveras, El Dorado & Mono, Placer, Tuolumne	0.987	0.778	1.073	534773	3.136
Marin, Sonoma	0.953	1.144	0.909	683315	2.365

<i>State adjusted difference in California</i>					
Counties	CB's	Alter. 1	Alter. 2	DD	II/DD
Imperial	4.269	4.928	4.085	125614	4.540
Kings	3.412	2.898	3.874	104834	4.291
Fresno, Madera, Mariposa	3.262	1.990	3.857	869960	4.884
Kern, Tulare	3.186	2.475	3.500	950411	4.550
Monterey, San Benito, Santa Cruz	2.995	2.496	3.244	652762	4.186
San Francisco	2.753	1.762	3.090	724551	4.475
Merced, Stanislaus	2.429	2.219	2.582	624309	3.668
Los Angeles	2.189	2.084	2.135	9014370	3.658
San Joaquin	2.060	1.417	2.259	523974	3.980
Sanluis Obispo, Santa Barbara	1.982	2.049	1.693	595458	3.087
Del Norte, Humboldt, Lake, Mendocino, Napa	1.968	1.970	1.913	393332	3.350
Inyo, San Bernardino	1.464	1.748	1.265	1629458	3.236
Riverside	1.344	1.369	1.360	1462999	3.283
Orange	1.276	0.944	1.491	2713751	3.323
San Mateo	1.228	0.866	1.351	674056	3.361
Santa Clara	1.137	1.076	1.219	1601952	3.179
Colusa, Glenn, Sutter, Tehama, Yolo	1.036	1.404	0.777	329003	2.780
Butte, Lassen, Modoc, Nevada, Plumas & Shasta, Sierra, Siskiyou, Trinity, Yuba	0.972	1.530	0.513	606664	2.491
Alameda	0.920	1.496	0.691	1376961	2.836
Alpine, Amador, Calaveras, El Dorado & Mono, Placer, Tuolumne	0.892	0.677	0.981	518002	3.238
Ventura	0.604	0.623	0.664	719791	2.806
San Diego	0.584	0.988	0.337	2645741	2.687
Sacramento	0.548	0.698	0.407	1165633	2.777
Contra Costa, Solano	0.112	0.672	-0.090	1285381	2.386
Marin, Sonoma	0.034	0.230	-0.011	667156	2.422

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