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Abortion access and state variation in observed unintended pregnancy

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Abstract

Objective—The state-level proportion of pregnancies that are unintended is an important social and public health indicator, and comparisons between states inform policy discussions. Unintended pregnancy is measured as a composite of abortions and unintended births, each of which is measured with error. We investigate whether between-state differences in abortion access and demand may bias comparisons between states' unintended pregnancy proportions when pregnancy intentions are misreported.

Study Design—We algebraically specify the model currently used to estimate unintended pregnancy, extend it to include underreporting, and simulate the impact of underreporting on observed unintended pregnancy. Comparing the impact of underreporting across states, we identify levels of underreporting at which between-state comparisons are compromised.

Results—We find that underreporting of unintended pregnancies could bias between-state comparisons when reporting of unintended pregnancies is less than 90-95%.

Conclusion—Current methods for estimating state-level unintended pregnancy proportions may underestimate unintended pregnancy to a greater degree in places with less abortion, and between-state comparisons may be biased. Estimates of state-level unintended pregnancy proportions would be more comparable if adjustment for completeness of retrospective underreporting were included in the estimation process.

Implications—Estimates of unintended pregnancy should be adjusted for non-sampling error and include variances based on sampling and non-sampling error in order to permit robust comparisons between states, between populations, and across time. More research on the fidelity of retrospective reporting of pregnancy intention would facilitate this endeavor.

Keywords

Unintended pregnancy; social indicators; reproductive health; abortion

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Introduction

State-level reproductive health policies have diverged, due in part to a deluge of state legislation affecting the funding of family planning, parental consent for minors, contraceptive access for undocumented migrants, postpartum contraceptive access, sexual education in the schools, and restrictions on the provision of abortion care [1]. Compounding this divergence, 15 states have yet to implement full Medicaid expansion, 28 have implemented it one way or another, and 7 states are still debating the issue [2]. In states that failed to expand Medicaid, some but not all provide access to family planning through Medicaid waiver programs, which are also targets of political attack [3]. In debates over the impact of these varied policies, an indicator that elected officials, policy advocates, and media are likely to call for and examine is state rate of unintended pregnancy.

Estimates of unintended pregnancy at the state level have only been available since 2011, when Finer and Kost compiled the necessary survey and vital registration data for each state and adapted the methods used to generate national-level estimates [4, 5].

Their state estimates have been used widely [6-9], but the relative standing of states did not correspond with states' levels of support for family planning and socioeconomic characteristics [10, 11]. To take one example, 53% of 2008 Texas pregnancies are estimated to be unintended, slightly lower than the 56% estimated for New York, which did not correspond with New York's more generous support for family planning [5, 12].

The question we explore is whether these anomalous rankings can be explained by underreporting. Unintended pregnancies that end in abortion are subject to little reporting error in many states because of vital registration, and in other states careful censuses have yielded reliable results [13, 14]. In contrast, the number of unintended live births is only known from retrospective surveys, which in the case of state-level estimates are mostly telephone interviews conducted a few months after delivery among a sample of women who gave birth during the year in question. In these interviews, respondents are asked whether just before getting pregnant they wanted to be pregnant sooner, later, at that time, or never again [15]. Demographers have been arguing for many decades over the reliability of answers to these questions, and the extent to which they are affected by *ex post* rationalization stemming from a reluctance to label a recently born child as mistimed or unwanted [16-20]. Without an independent accurate estimate of the level of intended fertility, the existence and magnitude of bias is difficult to assess, but there are indications that it may be substantial. Perhaps the most important is that many women report pregnancies that resulted from contraceptive failure as having been intended [21, 22]. A more direct indication of bias is that large discrepancies have been found when it has been possible to contrast prospective reports of intentions before a pregnancy with reports obtained after the pregnancy had been detected [16, 18, 22].

Given the uncertainty surrounding error and bias in the parameters used to estimate the intention status of live births in individual states, and in light of the fact that there are large differences between states in the proportion of all unintended pregnancies that depend on retrospective survey responses, we use simulation to test the sensitivity of between-state

comparisons in unintended pregnancy to different levels of completeness in the reporting of unintended live births.

We begin by algebraically specifying how Finer and Kost combine data from vital registration and surveys to estimate state unintended pregnancy [4]. We then extend their equations to include a term for the completeness of unintended pregnancy reporting. Finally, we perform a sensitivity analysis in order to calculate the level of underreporting at which between-state comparisons in unintended pregnancy are affected.

Background

Finer and Kost's procedures can be encapsulated in equations. These equations are the basis for our simulation model, and we derive them based on the authors' text descriptions. Note that we only follow Finer and Kost's estimation as far as the generation of the proportion of pregnancies that are unintended, and do not distinguish between pregnancies that are unwanted and those that are mistimed.

The proportion of pregnancies in a state that are unintended is the ratio of unintended pregnancies to all pregnancies. We represent this ratio as μ in the following equation:

$$\mu = \frac{A + B_U + M_U}{P} \quad (1)$$

where A is the number of abortions in the state (assumed for the purposes of this analysis to all be unintended), B_U is the number of unintended births in the state, M_U is the number of miscarriages of unintended pregnancies, and P is total pregnancies. The latter are the sum of all abortions (A), births (B), and miscarriages (M):

$$P = A + B + M \quad (2)$$

Thus, estimating μ requires estimates of the number of unintended pregnancies ending in birth (B_U), the number of unintended pregnancies ending in miscarriage (M_U), as well as the total number of births, abortions, and miscarriages. Only the total numbers of births and abortions are recorded in vital registration.

In order to estimate the remaining values, Finer and Kost first partition all pregnancies, P , into unintended pregnancies (P_U) and intended pregnancies (P_I) (see Figure 1). They then partition unintended pregnancies (P_U) into those the woman would end with abortion (P_A) and those the woman would carry to term (P_T) – were miscarriage not present. The pregnancies P_A resolve either through abortion (A) or miscarriage (M_A). Analogously, the pregnancies P_T resolve either through birth (B_U) or miscarriage (M_T).

There are three steps in the estimation of μ . The first generates an estimate of the number of unintended births, $\hat{B}_U$, as the product of the number of births from vital statistics (B), and the proportion of live births that are reported as unintended in the state, (p):

$$\hat{B}_U = B \cdot p \quad (3)$$

The quantity p is based on a retrospective question in each state's survey asking about pregnancy intentions, and is the proportion of survey respondents answering that they would have liked to have their child later or not at all.

The second step is inflation of all births and abortions to account for random miscarriage. Finer and Kost assume that 20% of all pregnancies that would have ended in birth and 10% of all pregnancies that would have ended in abortion end instead in miscarriage, as illustrated in Equation 4:

$$\hat{M} = 0.1 \cdot A + 0.2 \cdot B \quad (4)$$

The third step estimates the number of miscarriages of unintended pregnancies (M_U). Following the pregnancy partition scheme in Figure 1, the theoretical value of M_U is the sum of miscarriages among unintended pregnancies that would have ended in births (M_T) and miscarriages that would have ended in abortion (M_A). Following the assumption above regarding miscarriage among pregnancies that would have ended in either birth or abortion, the number of unintended pregnancies that are miscarried ($\hat{M}_U$) is a linear combination of abortions and unintended births.

$$\hat{M}_U = 0.1 \cdot A + 0.2 \cdot \hat{B}_U \quad (5)$$

Using their estimate of the number of unintended births ($\hat{B}_U$), together with the number of births and abortions, and their estimates of miscarriages by intention status, Finer and Kost construct their estimate of unintended pregnancy at the state level, $\hat{\mu}$, based on Equation 1 and the estimates generated in Equations 2 through 5.

$$\hat{\mu} = \frac{1.1 \cdot A + 1.2 \cdot B \cdot p}{1.1 \cdot A + 1.2 \cdot B} \quad (6)$$

One result we can derive from Equation 6 is the relative contribution to $\hat{\mu}$ from the second term in the numerator ($1.2 \cdot B \cdot p$). In this way, we can determine how states differ in their reliance on retrospective reporting.

Methods

Having specified the equations underlying Finer and Kost's calculations, we extend their model by introducing a parameter (B_U') for the completeness with which unintended pregnancies are reported retrospectively as unintended (c), and estimate the proportion of unintended births using this augmented model. Reporting is complete when $c=1$ and no unintended pregnancies ending in birth are reported retrospectively as unintended when $c=0$.

$$B_U' = B \cdot \frac{p}{c} \quad (7)$$

Note that this measure is identical to the unadjusted number of unintended live births, ($\hat{B}_U$), when reporting is perfect (when $c=1$).

Substituting B_U' for B_U in Equation 6, we define the resulting adjusted estimate of the proportion of pregnancies that are unintended as μ' :

$$\mu' = \frac{1.1 \cdot A + 1.2 \cdot B_U'}{1.1 \cdot A + 1.2 \cdot B} = \frac{1.1 \cdot A + 1.2 \cdot B \cdot \frac{p}{c}}{1.1 \cdot A + 1.2 \cdot B} \quad (8)$$

This adjusted estimate is only equivalent to Equation 6 when retrospective reporting is perfect (when $c=1$).

Based on Equations 6 and 8, we can write an expression for the difference between the adjusted and unadjusted estimate of the proportion of pregnancies that are unintended:

$$\mu' - \hat{\mu} = \left(1 - \frac{A}{\mu' \cdot P}\right) \cdot \frac{1}{1.1} \cdot (1 - c) \cdot \mu' \quad (9)$$

This difference is the error introduced by underreporting. The first factor, $\left(1 - \frac{A}{\mu' \cdot P}\right)$ is the proportion of all unintended pregnancies that do not end in abortion. The second factor deflates this number to address the absence of miscarriages in the data sources. The third factor $(1 - c)$, reflects the proportion of unintended pregnancies retrospectively reported as unintended. And the final factor (μ') is the adjusted proportion of all pregnancies that are unintended.

Equation 9 allows us to explore the responsiveness of observed unintended pregnancy ($\hat{\mu}$) to variation in the completeness of retrospective reporting (c). Because the equation has three unknown values, we choose pairs of states with plausibly similar levels of unintended pregnancy and plot the relationship between completeness of reporting (c) and observed unintended pregnancy ($\hat{\mu}$) assuming an identical and known value of the true underlying level of unintended pregnancy for each pair. We select Utah and New Hampshire because they are states with low unintended pregnancy but different levels of abortion. We select California and Texas for comparison as well, because both states are large, have similar demographic profiles, and have similar levels of unintended pregnancy but different levels of abortion.

We can also use Equation 9 to explore the relationship between the completeness of reporting (c) and the adjusted unintended pregnancy proportion, given states' observed unintended pregnancy proportion ($\hat{\mu}$), number of births (B), and number of abortions (A). We do this for a selection of states.

In Figure 2, we display a histogram of the variation across all states in the proportion of the unintended pregnancy estimate that relies on births for the year 2008, based on Finer and Kost's method and Kost's data [4, 5]. It shows substantial between-state variation in the proportion of unintended pregnancies measured via retrospective reporting. The proportion ranges from 98% in Wyoming to 34% in New York State. Forty-two of the fifty states fall along a relatively flat distribution ranging from 55% to 85%.

The data for each of the six states that we will use in our examples, as well the unadjusted estimate of the proportion of pregnancies that are unintended is shown in Table 1 [5]. These states were selected to represent a range of levels of abortion and a range of proportions of births reported as unintended. Since some states' administrative records for abortion are either incomplete or not available, in order to have consistent estimates of the total numbers of abortions in a state (A), we use the estimates available from the Guttmacher Institute census of abortion providers [5, 14, 23]. For consistency with previous results, we use the total number of births (B) and the proportion of births estimated to be unintended (p) from Kost's published tables for 2008 [5].

Results

In Figure 3, we use the equation for the sensitivity of measured unintended pregnancy to reporting (Equation 9) to illustrate the responsiveness of unintended pregnancy estimates to the completeness of reporting. We plot the trajectories of observed unintended pregnancy ($\hat{\mu}$) by completeness of reporting (c) using actual numbers of abortions (A) and pregnancies (P) for states (from Kost [5]). As described above, we picked two pairs of states with similar levels of unintended pregnancy but different levels of abortion to illustrate the differential impact of retrospective reporting as states' levels of abortion vary. In panel A, we plot the observed proportion ($\hat{\mu}$) by completeness of reporting (c) for California and Texas, assuming both states have adjusted unintended pregnancy proportions of 0.6. In panel B, we do the same for New Hampshire and Utah, assuming both states have adjusted unintended pregnancy proportions of 0.5. In both panels, the lines represent the positive relationship between completeness of retrospective reporting and estimated unintended pregnancy proportions. In general we found that states with higher levels of abortion are less responsive (have flatter slopes) to decreases in the completeness of retrospective reporting. The right hand side of each panel, the vertical line where reporting is complete ($c=1$), reflects the point where the estimated unintended pregnancy proportion is equal to the corrected unintended pregnancy proportion.

In Figure 4 we graph adjusted unintended pregnancy proportion ($\hat{\mu}$) by completeness of reporting (c) based on the equation for the adjusted proportion (Equation 8). This allows us to see how individual states' adjusted values increase as completeness of reporting (c) decreases. We do this for the six selected states. The points on the right hand side of the graphed lines (where the lines intersect the vertical line at $c = 1$) reflect the situation when retrospective reporting is perfect, and thus correspond to the proportions estimated by Finer and Kost ($\hat{\mu}$). The curves increase as completeness of reporting (c) decreases. Thus, as one might expect, less complete retrospective reporting (c) leads to a greater difference between the observed proportion ($\hat{\mu}$) and the adjusted proportion of pregnancies that are unintended

(μ'). In general we find that states with lower levels of abortion as a proportion of unintended pregnancies have steeper slopes and are thus more sensitive to changes in completeness of reporting (c). The exercise is based on the 2008 values (from Kost [5]) for the number of abortions (A), number of births (B), and the proportion of births reported as unintended (p).

Where the plotted lines in Figure 4 intersect, states' orderings by true unintended pregnancy reverse. For example, for values of c greater than about 0.9, New York has a higher unintended pregnancy proportion than California, whose own proportion is higher than the rate Texas. At levels of c less than about 0.9, however, Texas has a higher unintended pregnancy proportion than does either New York or California. Similarly, New Hampshire has a higher unintended pregnancy proportion than Utah at $c=1$, but at $c=.65$ the two lines cross and their ordering flips.

Figure 4 also allows us to see what the adjusted levels of unintended pregnancy (μ') would be if two states were to have different levels of reporting (c). For example, if reporting is low in Texas but high in New York, then the adjusted unintended pregnancy proportion in Texas may be much higher than it is in New York. Suppose, for example, that Texas' level of reporting is 80% ($c=0.8$) and New York's level of reporting is 95% ($c=0.95$). Then Texas' adjusted unintended pregnancy proportion is 61% and New York's is 55%. In this case, not only is their ordering switched, but Texas' unintended pregnancy proportion is substantially higher than New York's.

Discussion

By constructing estimates of unintended pregnancy at the state level, Finer and Kost [4] have substantially advanced our capacity to discuss unintended pregnancy in the United States. However, our modeling indicates that comparisons of US states' unintended pregnancy proportions are vulnerable to relatively small amounts of reporting error. Specifically, not taking reporting error into account may mean that estimates for states with low levels of abortion are disproportionately biased downward compared with estimates for states with higher levels of abortion.

The main limitation of our analysis is that it is hypothetical. We are unable to offer a reliable correction based on what is currently known about reporting bias. There is a need to know more about the reliability of the reporting of pregnancy intentions, and how it may vary across states and populations. For example, the stratification of reproduction by age, marital status, and race could lead women whose fertility is socially undervalued to report pregnancies as unintended [24]. Thus states with high levels of early, often non-marital pregnancy such as Mississippi may have more complete reporting of unintendedness. At the other extreme, women in communities with pronounced pronatalist values or who hold such values themselves might be reluctant to report a pregnancy as mistimed or unwanted. Thus, reporting may be less complete in states with large Latino populations [25], such as Texas or California, or where conservative religions are widely practiced and influential such as Utah and Idaho. State and national level studies that included prospective as well as retrospective assessment of pregnancy intentions could inform what we know about the fidelity of

reporting across groups and regions. Incorporating such findings into future estimates of unintended pregnancy at the state level could enhance the validity of between-state comparisons.

Another limitation of the model of unintended pregnancy presented here is that it is based on a dichotomous distinction between intended and unintended pregnancies. We have experimented with a more elaborate model that allows for a continuous scale of intendedness, and lets the probabilities of both abortion and reporting depend on the degree of intendedness [26, 27]. But while such a model might be more realistic, the results are essentially the same, and show a greater influence of reporting in states with lower rates of abortion.

In our exposition, we have focused on how comparisons of two states at a single point in time might be differentially affected by reporting error, but appraisals of trends in unintended pregnancy through time for a single state might also be affected by the differential utilization of abortion. In a state such as Texas where supply-side restrictions are having a large impact on abortion access [28], the proportion of pregnancies ending in abortion could decline substantially over a period of several years. The decline in abortion access in the state would surely increase the proportion of unintended pregnancies that end in births, but would have no effect on the underlying proportion of pregnancies that are unintended. However, the observed unintended pregnancy proportion could decrease as a function of the completeness of reporting. Similarly, some of the comparisons between groups shown in analyses of national survey data might be subject to the same bias. For example, one might question whether unintended pregnancy rates are truly lower among Evangelical Protestants than Mainline Protestants [29] since the former have a much lower proportion of pregnancies ending in abortion [8, 9].

Whether the level of aggregation is a state, a group, or even a nation, the measurement of unintended pregnancy should incorporate both sampling and reporting (non-sampling) error. Our algebraic specification is a starting point for a comprehensive specification of an estimate's uncertainty. This composite error would require probability distributions for the parameters c and p , the ratio of miscarriages to births and abortions, and A , the total number of abortions [30, 31]. Further research is needed to generate these distributions, which would permit robust comparisons between states, between populations, and across time.

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Figure 1.
Partition of pregnancies by intention and resolution

Figure 2.

States by proportion of reported unintended pregnancies ending in birth

Panel A. If both Texas and California had adjusted unintended pregnancy rates of 0.6

Panel B. If both New Hampshire and Utah had adjusted unintended pregnancy rates of 0.5

Figure 3.
Responsiveness of observed unintended pregnancy rate to variation in completeness of reporting

Figure 4.
Adjusted unintended pregnancy rate (μ') by completeness of reporting (c), selected states

Table 1

Parameter values for selected states

State	Abortions (A)	Live births (B)	Unintended pregnancy rate among live births (p)	Observed unintended pregnancy proportion ($\hat{\mu}$)
California	214,190	614,980	0.38	0.53
Mississippi	2,770	50,371	0.61	0.63
New Hampshire	3,200	13,733	0.27	0.40
New York	153,110	248,538	0.28	0.54
Texas	84,610	404,813	0.43	0.52
Utah	4,000	54,892	0.33	0.37

Values of A , B , and p calculated for 2008 based on [4,5]