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Research Article

Modeling the fertility impact of the proximate determinants: Time for a tune-up

John Bongaarts

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Modeling the fertility impact of the proximate determinants: Time for a tune-up

John Bongaarts[1]

Abstract

BACKGROUND
Many analyses of the determinants of fertility make a distinction between proximate and background determinants. The former include behavioral factors such as the use of contraception or abortion through which the background determinants (e.g., social and economic variables) affect fertility. These relationships were first recognized by Davis and Blake (1956), who defined a large set of "intermediate fertility variables." In the late 1970s Bongaarts (1978) identified a smaller set of proximate determinants and developed a relatively simple model to quantify their fertility effects.

OBJECTIVE
This paper fine-tunes the Bongaarts proximate determinants model in light of new evidence, research, and data that have become available over the past three decades. Reproductive behavior has changed substantially and certain original simplifying assumptions have become less accurate over time. In addition, new research allows some features of the model to be improved.

METHOD
Six adjustments to the model are proposed and implemented. The revised model is compared with the original version and with a revision proposed by Stover (1998).

RESULT
Revised estimates of the indexes of the proximate determinants and total fecundity are provided for the most recent DHS surveys in 36 developing countries. The revised model provides a better fit than do earlier models.

CONCLUSION
The proximate determinants model, as originally conceived, remains conceptually sound. However, theoretical and empirical evidence accumulated over the past three decades suggests a number of ways to fine-tune the model to make it more robust and accurate in contemporary populations. The resulting revised model provides an

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improved assessment of the roles of the proximate determinants in national and sub-national populations.

1. Introduction

The proximate determinants (PD) of fertility are the biological and behavioral factors through which the background determinants (social, economic, and environmental variables) affect fertility. The distinguishing feature of a proximate determinant is its direct connection to fertility. If a proximate determinant, such as contraceptive use, changes, then fertility necessarily changes also (assuming the other proximate determinants remain constant). This is not necessarily true for a background determinant of fertility such as income or education. Consequently, fertility differences among populations and trends in fertility over time can always be traced to variations in one or more of the proximate determinants. If accurately measured and modeled, the proximate determinants should explain 100% of variation in fertility.

These relationships were first recognized in the mid-1950s when Kingsley Davis and Judith Blake (1956) defined a large set of proximate determinants which they called the "intermediate fertility variables." This set was quite comprehensive and included some biological factors that differ little among populations[2]. In the late 1970s Bongaarts (1978, 1982) defined a somewhat different and smaller set of proximate determinants, thus simplifying the task of constructing models of human reproduction. His analysis indicated that four proximate determinants – marriage/cohabitation, contraception, induced abortion, and postpartum infecundability – are the most important for the analysis of fertility levels and trends. The identification of this smaller set of proximate determinants (PDs) led to the development of a relatively simple model that quantifies the fertility effect of each of these PDs.

The objective of this paper is to fine-tune this PD model in light of new evidence, research, and data that have become available over the past three decades. The model as originally conceived remains conceptually sound, and there is no reason to change the general multiplicative nature of the main equation. However, in recent decades reproductive behavior has changed substantially and certain original simplifying assumptions have become less accurate over time. In addition, new research allows some features of the model to be improved. A few revisions of features of the model are therefore desirable.

[2] Lactational amenorrhea was not included in the Davis Blake set of intermediate variables.

After a brief overview of the original PD model, the proposed revisions will be discussed and implemented. The revised model will then be applied to data from 36 developing countries using data collected in DHS surveys and compared with earlier models.

2. Background

At the core of the original PD model is the following multiplicative equation for a population at a given point in time

$$TFR = C_m C_c C_i C_a \, TF$$

where
TFR = Observed total fertility rate
C_m= Marriage index
C_c= Contraception index
C_i= Postpartum infecundability index
C_a= Abortion index
TF = Total fecundity rate.

The model treats each PD as a factor that inhibits fertility. Each index has values that range from 1 to 0 depending on the degree of inhibition. The index of marriage C_m measures the impact of the proportion of women in a marital union (i.e., formal marriage or consensual union). The index equals one when all women are cohabitating, and zero when no women are in a union. Only women in a union are assumed to be at risk of childbearing. The index C_c equals one when no contraception is used, and zero when all fecund women use 100% effective contraception. The index C_i equals one in the absence of lactational amenorrhea or postpartum abstinence and declines in size as the period of postpartum infecundability rises. The index of abortion equals one in the absence of abortion and declines as the incidence of abortion rises.

The total fecundity rate is the hypothetical total fertility rate that would be observed in a population in which all inhibiting effects of the proximate variables are absent, i.e., when $C_m = C_c = C_i = C_a = 1$. *TF* values typically are around 15 births per woman.

The equations for estimating the original values of the indexes are summarized in Table 1. Their calculation requires estimates of a number of variables, most of which can be readily obtained from DHS surveys. The main exception is the induced abortion rate, for which data are often lacking or of low quality.

Table 1: Original aggregate proximate determinants model and equations for indexes

	Equation	Variables
Original aggregate model	$TFR = C_m C_c C_i C_a \, TF$	*TFR*=total fertility rate *TF*= total fecundity rate
Marriage index	$C_m = \frac{\sum m(a) f_m(a)}{\sum f_m(a)}$	*m(a)*= proportion married by age f_m(a)= age-specific marital fertility rate *a*=age
Contraception index	$C_c = 1 - 1.08\, u\, e$	*u*=contraceptive prevalence (married women) *e*=average effectiveness
Postpartum infecundability index	$C_i = \frac{20}{18.5 + i}$	*i*=average duration of postpartum infecundability
Abortion index	$C_a = \frac{TFR}{TFR + b\, TAR}$	*TAR*= total abortion rate *b*=births averted per abortion
	$b = 0.4(1 + u)$	

Bongaarts and Potter (1983) also proposed an age-specific version of the PD model. It is summarized in Table 2. The age-specific equations for the indexes have the same structure as those in the aggregate model in Table 1. Age specific PD models have a clear advantage over the simple aggregate models because they take account of variation in the age structures of populations (Casterline, Singh, and Cleland 1984, Hobcraft and Little 1984, Singh, Casterline, and Cleland 1985). Unfortunately, applications of age-specific models have been very limited because they are more demanding of data.

The need for updating the original model has risen over time. Stover (1998) provides a detailed discussion of a number of these issues and proposed revisions to the equations for calculating indexes. Several of his ideas will be adopted below.

Table 2: Original age-specific proximate determinants model and equations for indexes

	Equations	Variables
Original age-specific model	$f(a) = C_m(a)C_c(a)C_i(a)C_a(a)\, f_f(a)$	$f(a)$ =age-specific fertility rate $f_f(a)$=age-specific fecundity rate *a*=age
Marriage index	$C_m(a) = m(a)$	*m(a)*= proportion married
Contraception index	$C_c(a) = 1 - r(a)u(a)e(a)$	*u(a)*=contraceptive prevalence (married women) *e(a)*=average effectiveness *r(a)*=fecundity adjustment
Postpartum infecundability index	$C_i(a) = \frac{20}{18.5 + i(a)}$	*i(a)*=average duration of postpartum infecundability
Abortion index	$C_a(a) = \frac{f(a)}{f(a) + b\, ab(a)}$ $b = 0.4(1 + u)$	*ab(a)*= abortion rate *b*=births averted per abortion

3. Proposed revisions

For each of the proximate determinants, specific issues have been identified as potentially requiring a revision. These issues and the solutions proposed here to address them are as follows:

3.1 Marriage/union/sexual exposure

Extramarital sex and pregnancy are becoming more prevalent in developed and developing countries (Caraël et al. 1995, MacQuarrie 2014; United Nations 2013). The assumption in the original PD model that sexual activity and childbearing only take place within marriages or consensual unions was always an issue in some populations, and has become increasingly less tenable. The solution proposed here is to estimate the number of women who are exposed to the risk of childbearing as the sum of married women (or in consensual unions as defined in DHS surveys) and unmarried women who are pregnant, report sex in the last month, use contraception, or are postpartum infecundable.[3] In addition the name of the index of marriage will be changed to the more accurate index of sexual exposure as proposed by Stover 1998.

3.2 Contraception

As the use of contraception has risen over time, the proportion of use that overlaps with postpartum infecundability has become significant in societies with long periods of breastfeeding or abstinence (Stover 1998). Ignoring this overlap (as was done in the original model) can therefore generate inaccurate results. This is particularly the case in countries with long durations of postpartum infecundability and with family planning programs that promote contraceptive use in the early postpartum months. The solution proposed here is to exclude overlap between contraceptive use and postpartum infecundabilty in the calculation of C_c.

A second issue related to contraception is that the original index of contraception is derived from the prevalence of contraception among all married women aged 15-49. This assumption implies that this prevalence (and hence C_c) is affected by the age distribution of married women. This is inconsistent with the other indexes, which are not affected by the age distribution of the population of women in unions. The solution proposed here is to use the age-specific PD model instead of the aggregate model.

A third issue related to contraception is that the original aggregate model takes account of variations in the average level of effectiveness of contraception as affected by the method mix, but it does not explicitly account for age differences in method effectiveness (Bongaarts and Potter 1983). In contrast, the original age-specific model takes account of variation in effectiveness by age but not by method. The solution proposed here is to revise the age-specific model by allowing variation in effectiveness by age and method (see further discussion in Appendix).

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[3] In the original model, Bongaarts recommended setting the marital fertility of age group 15–19 equal to 0.75 of the marital fertility of age group 20–24, because very erratic estimates for the age group 15-19 were observed in historical populations with low proportions married in this age group. With the inclusion of sexually active women there is no longer a need for such an adjustment, except in countries where the proportion sexually active in age group 15–19 is very small (e.g., less than 10%). In the set of 36 countries used in this study only one country (Rwanda) fell below this level and the marital fertility of age group 15–19 in this country is set equal to that of age group 20–24.

3.3 Postpartum infecundability

No revisions needed (except the model-related adjustment noted below)

3.4 Abortion

The abortion index is a function of the number of births averted by an abortion. In the original formulation, this number was estimated based on the level of contraceptive prevalence with an equation that had limited analytic foundation (Bongaarts and Potter 1993). Research by Bongaarts and Westoff (2000) has examined this issue, and a more accurate equation for the number of births averted per abortion is now available (see Appendix). This new equation will be used to calculate the abortion index in the revised model proposed here.

3.5 Model

The original model assumes that the proximate determinants at a point in time affect fertility at the same time. In reality, there is a nine-month delay between a change in a proximate determinant and its impact on fertility. In addition, the DHS surveys typically measure fertility for a three-year period before the survey (i.e., on average the *TFR* refers to 18 months before the survey). As a result there is a mismatch of 18+9 months between the timing of the measurement of the proximate determinants at the time of the survey and the *TFR*. This produces significant discrepancies in countries with rapidly rising contraceptive prevalence. To address this issue, the *TFR* based on births 3 years before the survey will be compared with PDs measured 27 months before the date of the survey in the applications presented below.

4. Comparison with revisions proposed by John Stover

The revisions to the original PD model proposed above overlap in a number of cases with those proposed by John Stover (1998). To simplify the discussion, the text below refers to the Stover revisions with the initials JS, and to the revisions proposed here by the present author, with the initials JB.

4.1 Marriage/union/sexual exposure

The two sets of revisions agree on the need to take into account sexual activity outside formal unions. However, the implementation of these revisions differs slightly. JS calculates the sexual activity index from the proportion of all women who have had sex in the last month or are pregnant or abstaining postpartum, but excludes married women who don't meet these criteria. In contrast, JB assumes that all women who are married or in a union are exposed to the risk of pregnancy, and includes all women (regardless of their marital status) who have had sex in the last month or are pregnant or abstaining postpartum or are contraceptive users. This slightly more inclusive definition of exposure is based on the assumption that most women who are in union and have sex less than once per month still should be considered at risk of pregnancy. It also seems plausible to assume that most contraceptive users are sexually active, even if not in the past month.

In several Asian and North African surveys, no information was collected from never-married women in the latest DHS. As a result, sexual exposure in these countries is underestimated to the extent that there is extramarital exposure. These countries will not be included in the empirical analysis below.

4.2 Contraception

The two sets of revisions agree on the need to address the overlap between postpartum amenorrhea and contraceptive use, but the implementation of these revisions also differs slightly. JS excludes overlap between contraceptive use and postpartum amenorrhea. JB excludes overlap between contraceptive use and postpartum infecundability, which may be due to either breastfeeding or abstinence.

Another significant JS revision involves the fecundity adjustment to the index of contraception. This issue is discussed below and in the Appendix.

4.3 Abortion

JS proposes a slight revision to the original equation for C_a to take account of contraceptive effectiveness. In contrast, the JB revision of C_a relies on the analysis of the impact of abortion by Bongaarts and Westoff (2000), which was published after Stover published his revisions in 1998.

4.4 Postpartum infecundability

JS and JB are in agreement, except that the latter takes into account the average 27-month delay between a change in postpartum infecundability and its impact on fertility.

4.5 Sterility

JS introduces a revised index of sterility C_f which is calculated as 1.0 minus the proportion of sexually active women estimated to be infecund. A woman is considered infecund at the time of a survey if she is not menopausal, postpartum amenorrheic, or pregnant; has not given birth in the last five years; has not used a contraceptive and has been in a union during that period. Women who declare themselves to be infecund are also included. This is the approach used by the DHS.

In contrast, JB does not discuss the sterility index because there appears to be no need for it in contemporary populations. Introducing an index of sterility would be useful if variations in sterility exist among populations and contribute significantly to variations in fertility. There is no doubt that sterility was at elevated levels in selected African populations in the decades before the 1980s (Frank 1983). Beginning in the 1990s, however, the role of sterility variations appears to have become small enough to be ignored. This conclusion is based on the lack of variation in the proportion of childlessness reported by women aged 40–49 in recent DHS surveys. In the absence of pathological sterility, around 3 percent of women in a union are childless at the end of the childbearing years. In the 67 countries with DHS surveys conducted since 2000, the average proportion childless is 2.2% and in only one country (Jordan) does this proportion exceed 5% (the reason for this outlier is unlikely to be pathological sterility, because this proportion in Jordan was 3% in the 1990s). These statistics suggest no significant variation in pathological sterility in recent surveys. An examination of trends shows that the proportion who were childless was slightly elevated in a few African countries in the 1990s, but it subsequently declined to low levels. For example, in Cameroon, childlessness among women aged 40-49 declined from 9.9% in the 1991 survey to 2.4% in 2011. In interpreting these trends it is important to note that these proportions at the end of the childbearing years reflect sterility risks in earlier decades when the women were in their peak childbearing years. Therefore, the 9.9% estimate for Cameroon in 1991 mostly reflects elevated risks of sterility between the late 1960s and early 1980s. The low sterility levels measured two decades later in the 2011 survey suggest that sterility declined to low levels starting in the 1980s. These findings indicate no significant elevated sterility levels among women in their childbearing years in the

1990s and later[4]. There seems, therefore, to be no need to introduce a sterility index to help explain variation in fertility in contemporary populations.

It should be noted that the sterility index proposed by JS shows surprising and unexplained variation among populations. For example, the proportion of sexually active women who are infecund in Asian countries is much higher than in Latin American countries (Stover 1998). There is no direct evidence that pathological sterility is higher in Asia than in Latin America. It is more plausible to assume that these differences are attributable to errors in the DHS calculations of infecundity or to the inaccurate reporting of several of the survey questions needed to estimate infecundity.

In sum, JS and JB are in broad agreement on a number issues – the revisions to include extramarital sexual exposure and to exclude overlap between contraceptive use and postpartum infecundability – even though implementations differ somewhat. But they have differing views on the usefulness of a sterility index. In addition, JB includes model changes not proposed by JS, such as the age weighting of contraceptive prevalence in the index of contraception, which is necessary to make the model analytically sound.

5. Implementation of revisions

Table 3 presents the revised equations for the age-specific PD model. The calculation of these indexes requires the following age-specific variables:

$m(a)$= proportion married/in union
$ex(a)$= extramarital sexual exposure
$u(a)$= contraceptive prevalence (among sexually active women)
$o(a)$= contraceptive prevalence that overlaps with postpartum infecundability
$e(a)$= average contraceptive effectiveness
$r(a)$= fecundity adjustment
$i(a)$= average duration of postpartum infecundability
$f(a)$= fertility rate
$f_m(a)=f(a)/(m(a)+ex(a))$= fertility rate among sexually exposed women
$ab(a)$= abortion rate

[4] Although HIV positive women have elevated rates of spontaneous abortion (Gray et al. 1998), the DHS estimates of childlessness indicate no significant effect of the HIV epidemic on permanent sterility. In addition, the epidemic appears to have had little net impact on total fertility rates (Fortson 2009)

All but two of these variables can be estimated from DHS surveys using procedures described by Bongaarts and Potter (1983). The exceptions are *r(a),* and *ab(a),* which require some further discussion, provided in the Appendix.

Table 3: Revised age-specific proximate determinants model and equations for indexes

Index	Equations	Variables
	$f(a) = C_m^*(a)C_c^*(a)C_i^*(a)C_a^*(a)\ f_f^*(a)$	*****represents revised measures
Sexual exposure index	$C_m^*(a) = m(a)+ex(a)$	*m(a)*= proportion married/union *ex(a)*=extramarital exposure
Contraception index	$C_c^*(a) = 1 - r^*(a)(u^*(a) - o(a))e^*(a)$	u^* *(a)*=contraceptive prevalence (exposed women) *o(a)*=overlap with postpartum infecundability e^**(a)*=average effectiveness r^**(a)*=fecundity adjustment
Postpartum infecundability index	$C_i^*(a) = \frac{20}{18.5 + i(a)}$	*i(a)*=average duration of postpartum infecundability
Abortion index	$C_a^*(a) = \frac{f(a)}{f(a) + b^*\ ab(a)}$ $b^* = \frac{14}{18.5 + i(a)}$	*ab(a)*= abortion rate

*represents revised measures

The revised aggregate model is summarized in Table 4. The aggregate indexes are weighted versions of the age-specific indexes, with the weights varying by index. As a result, the aggregate and age-specific models are completely consistent with one another.

Table 4: Revised aggregate proximate determinants model and equations for indexes

Index	Equations	Variables
Revised aggregate model	$TFR = \sum C_m^*(a) C_c^*(a) C_i^*(a) C_a^*(a)\, f_f^*(a)$ $= C_m^* C_c^* C_i^* C_a^*\, TF^*$	TF^* = revised total fecundity rate $f_f^*(a)$ =revised fecundity rate
Sexual exposure index	$C_m^* = \sum C_m^*(a)\, w_m(a)$ $w_m(a) = \frac{f_m^*(a)}{\sum f_m^*(a)}$ $f_m^*(a) = C_c^*(a) C_i^*(a) C_a^*(a)\, f_f^*(a)$	$f_m^*(a)$= fertility rate, exposed women *a*=age
Contraception index	$C_c^* = \sum C_c^*(a)\ w_c(a)$ $w_c(a) = \frac{f_n^*(a)}{\sum f_n^*(a)} \approx \frac{f_f^*(a)}{\sum f_f^*(a)}$ $f_n(a) = C_i^*(a) C_a^*(a)\, f_f^*(a)$	$f_n^*(a)$= natural exposed fertility a=age
Postpartum infecundability index	$C_i^* = \sum C_i^*(a)\, w_i(a)\ \approx\ C_i$	
Abortion index	$C_a^* = \sum C_a^*(a) w_a(a)\ \approx \frac{TFR}{TFR + b^*\, TAR}$	

*represents revised measures

6. Results

Estimates of each of the revised indexes were obtained from data collected in the latest available DHS surveys in 36 countries. Countries are limited to those that have at least two standard all-women DHS surveys (to allow interpolation of the proximate determinants) and ex-Soviet countries are excluded.[5] Given the significant revisions proposed here, it is of interest to briefly examine the differences between the three models being examined here (i.e., the original version and the revisions by JS and JB).

Figures 1 to 4 plot the JS and JB indexes versus the original indexes. As expected, the three sets of indexes are positively correlated. The strongest correlations are observed for C_c, C_i, and C_a, and the weakest for C_m. The latter finding is in part due to an outlier, Namibia, where extramarital sex is relatively common. Namibia's original

[5] Eritrea 2000 is also excluded, due to the fertility-depressing effect of the war with Ethiopia (Blanc 2004).

value for C_m equaled 0.30 while the JC and JB revised values for C_m equal 0.51 and 0.74 respectively.

Figure 1: Sexual exposure index, original vs. revised

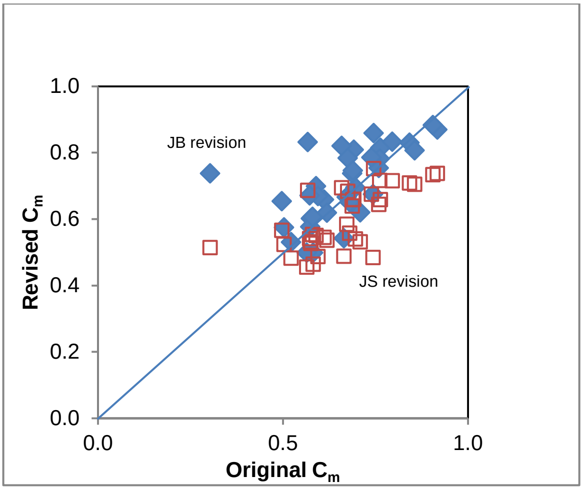

Figure 2: Contraceptive index, original vs. revised

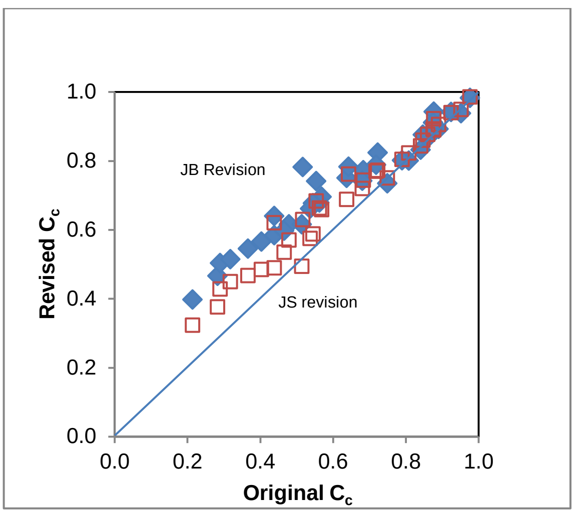

Figure 3: Postpartum index, original vs. revised

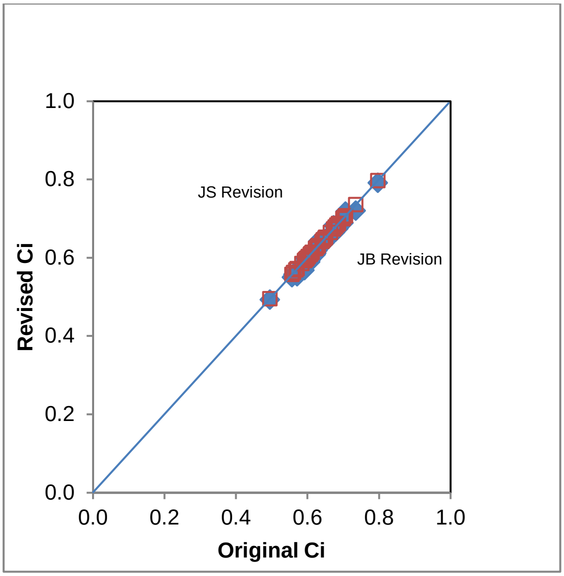

Figure 4: Abortion index, original vs. revised

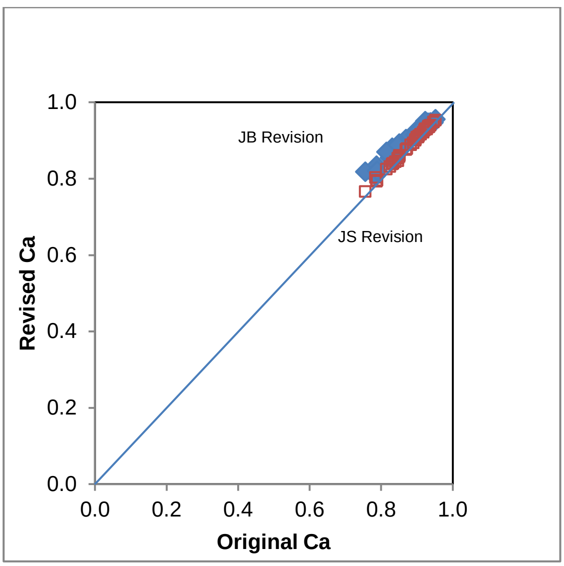

Table 5 presents the unweighted averages for the four indexes for the different models. The three models provide quite similar estimates for C_i and C_a, but differences are substantial for C_m and C_c. Particularly notable is the difference in C_m between the JS and the JB models (0.60 vs 0.70) which is primarily due to the way women who are married but did not have sex in the last month are treated. This group of women is excluded in the JS model but is included in the JB model. Differences in average values of C_c are also substantial. This is in part due to the exclusion of contraceptive overlap with postpartum infecundability (which raises C_c) in the JS and JB models and to the use of age-weighing (which also raises C_c) in the JB model[6].

Table 5: Unweighted average of indexes for 36 countries

	C_m	C_c	C_i	C_a
Original model	0.66	0.63	0.63	0.88
JS revision	0.60	0.70	0.63	0.88
JB revision	0.70	0.73	0.62	0.90

Figures 5 to 8 plot the three sets of indexes by the *TFR*. As expected, the indexes C_m, C_c, and C_a decline as countries move from high to low fertility and the inhibiting effects of these PDs become stronger. In contrast, C_i rises as countries move through the transition because breastfeeding and postpartum abstinence decline. The largest differences between models are observed for C_m and C_c as expected from the findings in Table 5. Interestingly, the slopes of the regression lines fitted to the three model estimates are similar for C_m. In contrast, the slopes for C_c vary by model, and countries with low fertility have the largest differences between models.

Estimates of the JB revised indexes as well as *TFR* and *TF* for individual countries are provided in Table 6. Note that the average value of *TF* for the 36 countries is 15.4, which is close to the estimate of 15.3 provided by Bongaarts and Potter (1983).

[6] The JS C_c also differs from the original model because it excludes a fecundity adjustment.

Figure 5: **C_m by TFR**

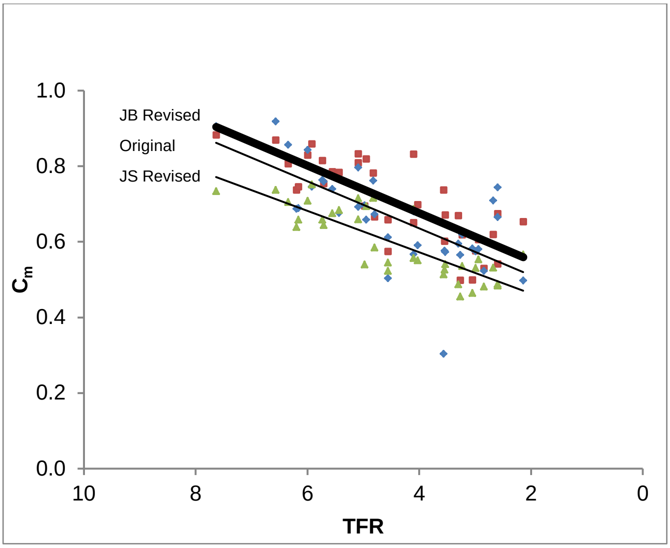

Figure 6: **C_c by TFR**

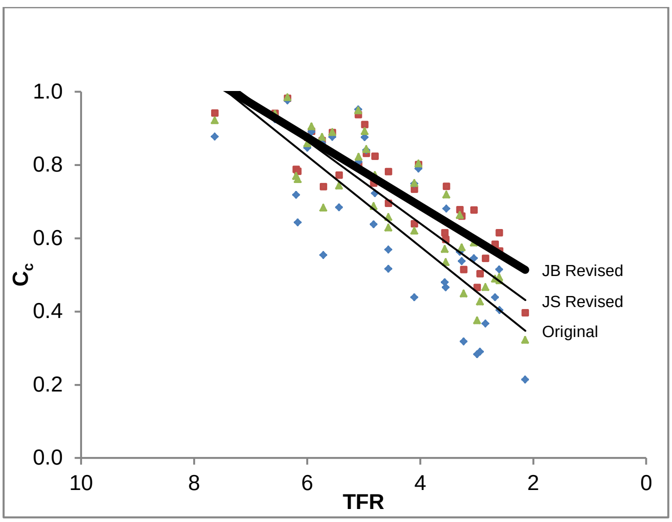

Figure 7: C_i by TFR

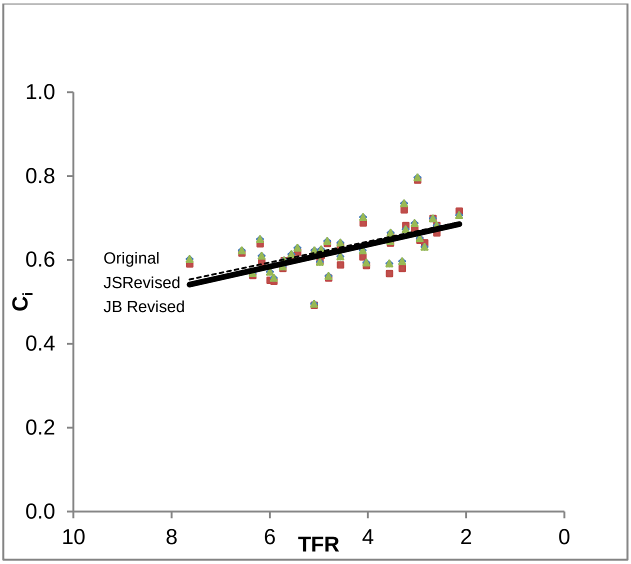

Figure 8: C_a by TFR

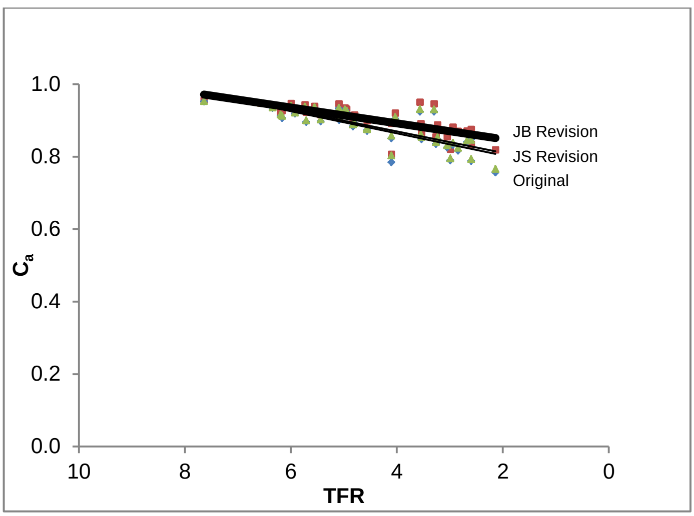

Table 6: Estimates of proximate determinant indexes and TFR, TF, and TFR_e for 43 DHS countries

	TFR	C_m	C_c	C_i	C_a	*TF*	*TFRe*
Benin 2006	5.7	0.81	0.87	0.58	0.94	14.8	5.9
Burkina Faso 2010	6.0	0.83	0.88	0.55	0.95	15.8	5.8
Cameroon 2011	5.1	0.81	0.80	0.62	0.92	13.9	5.6
Chad 2004	6.3	0.81	0.98	0.56	0.94	15.2	6.4
Cote d'Ivoire 2011-12	5.0	0.82	0.83	0.61	0.93	12.8	6.0
Ethiopia 2011	4.8	0.67	0.82	0.56	0.91	17.2	4.3
Gabon 2012	4.1	0.83	0.73	0.69	0.81	12.1	5.2
Ghana 2008	4.0	0.70	0.80	0.59	0.92	13.3	4.6
Guinea 2012	5.1	0.83	0.94	0.49	0.95	14.0	5.6
Kenya 2008-09	4.6	0.66	0.70	0.63	0.90	17.5	4.0
Lesotho 2009	3.3	0.67	0.68	0.58	0.95	13.3	3.8
Madagascar 2008-09	4.8	0.78	0.75	0.64	0.90	14.2	5.2
Malawi 2010	5.7	0.75	0.74	0.60	0.92	18.5	4.7
Mali 2006	6.6	0.87	0.94	0.62	0.95	13.8	7.3
Mozambique 2011	5.9	0.86	0.89	0.55	0.93	15.1	6.0
Namibia 2006-07	3.6	0.74	0.62	0.57	0.95	14.6	3.8
Niger 2012	7.6	0.88	0.94	0.59	0.96	16.2	7.2
Nigeria 2008	5.6	0.79	0.89	0.61	0.94	14.0	6.1
Rwanda 2010	4.6	0.57	0.78	0.59	0.90	19.1	3.7
Senegal 2010-11	5.0	0.69	0.91	0.60	0.93	14.1	5.4
Tanzania 2010	5.4	0.78	0.77	0.62	0.92	15.8	5.3
Uganda 2011	6.2	0.74	0.79	0.64	0.92	18.1	5.3
Zambia 2007	6.2	0.75	0.78	0.59	0.93	19.2	4.9
Zimbabwe 2010-11	4.1	0.65	0.64	0.61	0.89	18.2	3.5
Cambodia 2010	3.1	0.50	0.68	0.67	0.85	15.7	3.0
India 2005-06	2.7	0.62	0.58	0.70	0.87	12.2	3.4
Indonesia 2012	2.6	0.54	0.57	0.68	0.83	14.9	2.7
Nepal 2011	2.6	0.67	0.61	0.66	0.88	10.8	3.7
Philippines 2008	3.3	0.50	0.66	0.72	0.85	16.1	3.1
Bolivia 2008	3.5	0.60	0.60	0.64	0.89	17.3	3.1
Colombia 2010	2.1	0.65	0.40	0.72	0.82	14.1	2.3
Dominican Rep. 2002	3.0	0.58	0.47	0.79	0.82	17.2	2.7
Haiti 2012	3.5	0.67	0.74	0.66	0.87	12.4	4.4
Honduras 2011-12	2.9	0.61	0.50	0.65	0.88	16.9	2.7
Nicaragua 2001	3.2	0.62	0.51	0.68	0.89	16.8	3.0
Peru 2000	2.8	0.53	0.54	0.64	0.87	17.7	2.5
All surveys	4.46	0.70	0.73	0.62	0.90	15.4	4.50

7. Comparing the accuracy of models

A common approach to assess the accuracy of PD models is to compare the observed TFR with a model estimated TFR_e in a set of countries (Bongaarts and Potter 1983, Stover 1998). The value of TFR_e in a country is obtained by a three step process: 1) calculate the total fecundity rate (*TF*) by dividing the observed *TFR* by the product of all indices, 2) average the values of *TF*s in a group of countries (36 in the present study) to obtain TF_{ave} and 3) multiply TF_{ave} by the product of all country specific indices to obtain TFR_e. That is, for the original and JB models

$$TFR_e = C_m^* C_c^* C_i^* C_a^* \; TF_{ave}$$

The JS model has a similar equation, but differs slightly because it has different subscripts for the indexes, and it includes a fifth index of sterility[7] (see Stover 1998 for details). Despite these differences, the process for estimating TFR_e is the same.

The results of this comparison are summarized in Table 7. The first and second rows of this table present the unweighted averages of the observed and model-estimated total fertility rates for the 36 countries. The average error in the estimates (i.e., TFR_e-*TFR*) in the third row gives the bias for the models. A perfect model would have no bias. Instead, the results show that all three models have a positive bias ranging from 1.19 (births per woman) for the original model to 0.04 for the JB model. A second indicator of the accuracy of the model is the standard deviation of the error which is presented in the last row of Table 7. In a perfect model, this standard deviation would be zero. In the populations examined here, the standard deviation of the error is positive with the largest value (1.47) for the original model and the lowest (0.61) for the JB model. It should be noted that both the JS and JB revised models perform much better than the original model, in part due to a few outliers such as Namibia. In addition, the JB model is somewhat more accurate than the JS model according to the measures used in Table 6.

Table 7: Unweighted average of observed and model estimates of total fertility rates for 36 countries and mean and standard deviation of error

	Original (Bongaarts)	**JS revision (Stover)**	**JB revision (Bongaarts)**
Observed TFR	4.46	4.46	4.46
Model estimated *TFRe*	5.65	4.64	4.50
Error (TFR_e- *TFR*)	1.19	0.18	0.04
Standard deviation of error	1.47	0.76	0.61

[7] In addition, the level of fertility in the absence of inhibiting factors is called potential fertility.

But none of the models is perfect. There are several possible explanations for the differences between *TFR* and TFR_e in all three models:

1) Errors in the estimates of the *TFR* (both sampling and non-sampling)
2) Errors in measures of the proximate determinants
3) Errors in the model equation for estimating the indexes and *TF*
4) True variation in *TF* due to differences in frequency of intercourse and sterility.

Unfortunately, it is not possible to quantify the roles of these potential errors.

8. Conclusion

The original proximate determinants model developed in the late 1970s by Bongaarts has found wide use in the analysis of levels, differentials, and trends in fertility. Theoretical and empirical evidence accumulated over the past three decades suggests a number of ways to fine-tune the model to make it more robust and accurate in contemporary populations. Six adjustments are proposed and implemented in this study. Compared with the original model and the revisions proposed by Stover (1998), the resulting new revised model provides an improved assessment of the roles of the proximate determinants, as indicated by a greater accuracy in predicting the TFR in a group of 36 countries with recent DHS surveys. The aggregate version of the revised model is directly derived from the age-specific model, which provides a sounder analytic basis for estimating the fertility impact of the proximate determinants. This is particularly the case for applications to sub-national populations (e.g., level of education or urban-rural) because differences in age structures can be substantial.

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Appendix: Procedures for estimating variables and parameters

1. Births averted per abortion

An analysis of the tradeoff between abortion and contraception by Bongaarts and Westoff (2000) estimates the number of births averted per abortion as the ratio of the average reproductive time associated with an abortion to the average reproductive time associated with a live birth. Following Bongaarts and Potter (1983), the former can be estimated as 14 months and the latter as 18.5 plus the average postpartum infecundability interval. The revised equation for the number of births averted per abortion therefore is

$$b^* = \frac{14}{18.5+i} \qquad \text{(A1)}$$

In theory, this equation should allow for variation in i by age in the age-specific version of the revised model. But in practice the postpartum infecundability interval rises only slightly by age, and the other components of birth and abortion intervals also rise as the waiting time to conception increases with age. The net result is only a small variation in $b*$ by age, and in the applications in this study $b*$ and i will be given the average value at all ages.

2. Age-specific abortion rates

In developed countries with reasonably accurate abortion statistics, abortion rates by age have an inverted U graph-shape, with peak rates between ages 20 and 29 (Sedgh et al. 2012a). For present purposes it will be assumed that the countries included in this paper also have an inverted U graph-shape. Given the absence of accurate age data for abortion rates, it will further be assumed that this shape (but not the level) is the same as that of the age-specific fertility rates. With this simplifying assumption, the ratio of age-specific abortion rates to age-specific fertility rates is equal to *TAR/TFR* at all ages. As a result, $C_a^*(a) = C_a^*$.

Estimates of *TAR* are calculated as 30 times the abortion rate per 1000 women aged 15–45 (divided by 1000). The abortion rate for individual countries is obtained by assigning each country the abortion rate of its world sub-region as estimated by Sedgh et al. (2012b). The exception is Gabon, for which country estimates are available from other sources.[8]

3. Fecundity adjustment factor *r(a)*

The parameter *r(a)* adjusts the prevalence of contraceptive use among all exposed women to account for the fact that prevalence is higher among fecund women than among infecund women. To estimate this parameter, a slightly revised version of a method developed by Bongaarts and Kirmeyer (1982) is applied. The method starts by introducing a new variable $f_{nc}(a)$ which is defined as the fertility rate of exposed women that would be observed in the absence of abortion and postpartum infecundability. As a result

$$f_{nc}(a) = \frac{f(a)}{C_m^*(a)C_i^*(a)C_a^*(a)} \qquad \text{(A2)}$$

$$= f_f^*(a)C_c^*(a) \qquad \text{(A3)}$$

$$= f_f^*(a)\,[1 - r^*(a)(u^*(a) - o(a))e^*(a)] \qquad \text{(A4)}$$

In this equation $f_{nc}(a)$ can be calculated with equation (A2) from estimates of $f(a)$, $C_m^*(a), C_i^*(a),$ and $C_a^*(a)$ using data from DHS surveys. The product $(u^*(a) - o(a))e^*(a)$ can also be estimated from DHS data on contraceptive prevalence and method mix by age (and using standard estimates of effectiveness by method from Bongaarts and Potter (1983)).

The unknowns in equation (A3) are the fecundity rate $f_f^*(a)$ and the fecundity adjustment factor $r^*(a)$. These unknowns are estimated with a set of OLS regressions (one for each age group) as proposed by Bongaarts and Kirmeyer (1982). In each regression $f_{nc}(a)$ is the dependent variable and the product $(u^*(a) - o(a))e^*(a)$ is the independent variable. These regressions yield estimates of the intercept (which estimates $\bar{f}_f(a)$, the average of $f_f^*(a)$ for all countries) and the slope (which estimates the average impact *S(a)* of a unit increase in $(u^*(a) - o(a))e^*(a)$ on $f_{nc}(a)$ for all countries. From these regressions $r^*(a)$ is estimated as $S(a) = r(a)f_n(a)$. The regression results are shown in table A1.

[8] For Gabon, reported abortion rates were taken from the DHS first country report and inflated by 150% on the assumption that underreporting of abortion is the same as in the US (Institute of Medicine 1995**).**

Table A1: Results of OLS regression of $f_{nc}(a)$ on $(u^*(a) - o(a))e^*(a)$

	15–19	20–24	25–29	30–34	35–39	40–44	45–49
Intercept $\bar{f}_f(a)$	679	631	588	514	380	192	60
Slope S(a)	-416	-516	-586	-546	-409	-224	-88
$r^*(a) = S(a)/\bar{f}_f(a)$	0.62	0.81	0.99	1.08	1.14	1.26	1.62

The above values of $\bar{f}_f(a)$ and $r^*(a)$ are used in the calculations of the model index $C_c^*(a)$.

Note that the fecundity adjustment $r^*(a)$ is lower than one in age group 15-19 and 20-24. This finding is explained by the fact that *e(a)* as calculated from the method mix does not include the effect of age on method effectiveness (e.g. because method effectiveness among spacers is lower than among limiters). The above estimate of $r^*(a)$ therefore also picks up the effects of age on method effectiveness. Since the youngest age groups have below-average method effectiveness, the value of $r^*(a)$ falls below 1. At higher ages the shape of the $r^*(a)$ function is similar to the one estimated by Bongaarts and Kirmeyer (1982).

To check the plausibility of the age patterns of the indexes, estimates of age-specific fecundity rates were calculated for all countries with

$$f_f^*(a) = \frac{f(a)}{C_m^*(a)C_c^*(a)C_i^*(a)C_a^*(a)}$$

The resulting age patterns of individual countries were similar in shape to $\bar{f}_f(a)$. The only clear anomalies occurred in age group 45-49 in Colombia, Dominican Republic, and India, where $f_f^*(a)$ values were slightly negative. It is likely that the direct cause of this finding in these cases is a negative value of $C_c^*(a)$. These countries have high levels of sterilization, and the standard value of $r^*(a)$ is apparently too high. It would be possible to make $r^*(a)$ a function of the proportion sterilized, but that would introduce complications which add little to the accuracy of the aggregate results, which are insensitive to errors in the oldest age groups where fertility is very low. Instead, a simpler solution is used here: any value of $C_c^*(a)$ that is negative or less than 0.1 is set equal to 0.1. With this adjustment, the above anomalies disappear and all the country patterns $f_f^*(a)$ are broadly similar in shape to $\bar{f}_f(a)$.