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Trends in US sex ratio by plurality, gestational age and race/ethnicity†

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BACKGROUND: The sex ratio in the USA has declined over recent decades, resulting in fewer male births. Concurrent changes in the childbearing population may have influenced the sex ratio, including increases in multiple births, improvements in perinatal survival and increased Hispanic births.

METHODS: Data from the US natality files (1981–2006) were analyzed to determine the impact of changes in birth characteristics on male birth proportion. Male birth proportion was calculated as the number of male births divided by the total number. In separate analyses, trends in male birth proportion from 1981 to 2006 were adjusted for plurality (singleton, multiple), gestational age (<28, 28–32, 33–36, ≥37 weeks) and, from 1989, maternal Hispanic ethnicity. Separate analyses were conducted for white and black births. Log binomial regression was performed to estimate crude and adjusted trends with year as independent variable.

RESULTS: Trends in male birth proportion differed significantly according to plurality among white ($P < 0.01$), but not black births. Adjustment for gestational age tempered the trends among white singletons ($P < 0.0001$) and multiples ($P < 0.05$) but had no effect on trends in black male birth proportion. Adjustment for Hispanic ethnicity had no impact on trends in black male birth proportion and any effect on white births was negated by changes in gestational age trends.

CONCLUSIONS: Lack of consistent influences on, or patterns of change in, the proportion of male births between different subpopulations of births suggests that a single mechanism is unlikely to explain the oft-referenced decrease in the overall US sex ratio.

Key words: sex ratio / statistics / epidemiology / USA / ethnicity

Introduction

The sex ratio, or the ratio of male to female births, is used in demographic and environmental studies to assess changes that may indicate an imbalance in fertility or birth events within a population. The US sex ratio has decreased slightly over the last few decades (Mathews and Hamilton, 2005) and this decrease has given rise to speculation regarding its cause.

The lower-than-expected proportion of males born to fathers exposed to dioxin highlights the possible influence of industrial contaminant exposure on a population's sex ratio (Mocarelli *et al.*, 2000). Thus, potential increases in low-level exposure to hormonally-active compounds have been postulated to underlie the recent US sex ratio trends in some analyses (Davis *et al.*, 2007). Other analyses, based on similar data, have highlighted the inconsistent trends in the sex ratio among births to white women (generally decreasing) and black women (generally increasing) as evidence of natural variation (Maconochie and Roman, 1997; Marcus *et al.*, 1998; Nicolich *et al.*, 2000; Mathews and Hamilton, 2005); a more consistent effect might be expected if pervasive environmental exposures influence the sex ratio at birth within the USA.

Over the last several decades, other changes in the US childbearing population may have plausibly influenced the sex ratio. In the last 20 years there have been unprecedented increases in births to older mothers, multiple births (twins, triplets, and higher order births) and Hispanic births. At the same time great strides have been made in improving perinatal survival and decreasing fetal deaths, particularly among infants born at earlier gestational ages, leading to an increasing proportion of live births at earlier gestational ages. Each of these factors is potentially associated with a changing sex ratio and as this involves a fairly large portion of US births, it is plausible that temporal trends in these factors may influence the overall trend in the US sex ratio at birth.

The objective of this study is to assess trends in the sex ratio at birth among US subpopulations characterized by recent changes that may

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have influenced the overall US sex ratio, specifically: (1) multiple births—the sex ratio among multiple births is lower than that among singleton births (James, 1975) and evidence suggests an association between ovarian stimulation and decreased male proportion (Sampson *et al*., 1983; Dickey *et al*., 1995); (2) changes in perinatal care and fetal survival—the influence of the interactions among a decreasing fetal death rate (MacDorman *et al*., 2007), increasingly higher male proportion among fetal deaths (Davis *et al*., 2007), and increasing trends for early preterm birth (Branum and Schoendorf, 2002) on overall sex ratio is difficult to predict given the lack of understanding of the mechanisms underlying those trends; and (3) changes in the percentage of births to Hispanic mothers—a relatively low proportion of male births has been reported among infants born in California to Mexican mothers (Smith and von Behren, 2005), although not to the same degree as among the overall Hispanic population in the USA (Mathews and Hamilton, 2005). Given the large increase in the percentage of births to Hispanic women, the potential influence of these births on the overall sex ratio warrants evaluation. Secondarily, we examined the influence of maternal age on these factors because of its known associations with plurality, preterm birth and sex ratio, and the increase in maternal age at childbirth over recent years. Understanding the patterns of sex ratio changes in the USA may aid in the interpretation of potential causes underlying the recent decrease in the US sex ratio at birth.

Materials and Methods

Data

To examine overall trends and the role of plurality on observed trends, we analyzed US Centers for Disease Control and Prevention/National Center for Health Statistics (CDC/NCHS) natality data files from 1981 to 2006. Prior to 1978, many US states did not report gestational age on the birth certificate; therefore, there is a high percentage of missing data for 1972–1978, relative to the other years. Goodness-of-fit tests also indicated poor data quality prior to 1981. The end time points of analysis represent the most recent data on births available at the time of analysis. For each year included in these analyses, there are ∼3–4 million births.

Natality files were restricted to records with the following criteria regarding gestational age reporting. First, only records with plausible birthweight-gestational age combinations, according to a previously described algorithm (Alexander *et al*., 1996), were included in analysis. On average, implausible birthweight/gestational records comprised 0.5% of the data and this varied little over time. Between 1981 and 1988, only measures of gestational age based on last menstrual period (LMP) were available on the birth certificate. In 1989, a clinical estimate of gestational age was introduced, accounting for approximately 5% of data from 1989 to 2006. However, use of this estimate varies by US state and possibly demographic characteristics. In order to reduce possible bias, only gestational ages based on LMP were included in analysis. Finally, any birth record with missing LMP-based gestational age was excluded.

To examine the influence of increasing numbers of births to mothers with Hispanic ethnicity on trends, less data were available (1989–2006). In 1989, the Federal government began using the classification of Hispanic ethnicity to differentiate race from ethnicity in natality statistics. Hispanic ethnicity was classified based on reports of Mexican, Puerto Rican, Cuban, Central and South American heritage. Persons not of Hispanic ethnicity who report themselves as black or white are categorized as non-Hispanic black or non-Hispanic white, respectively (Macdorman *et al*., 2007). Hispanic ethnicity was not reported by three states in 1989 (NH, OK, LA), two states in 1990 (NH, OK) and one state in 1991 and 1992 (NH): this accounted for 3.4% of all births in 1989, 1.5% of all births in 1990 and 0.4% in 1991. Data from these states were excluded from analysis for the corresponding years. In addition there were <1% of births with missing Hispanic ethnicity among black and white births over time. These records were also excluded from analysis.

Variable definitions

For purposes of this analysis, the proportion of male births out of all births, or group-specific births, serves as the dependent variable. For the primary analyses, maternal race was coded as White or Black. Other race groups were excluded. To examine the role of Hispanic ethnicity on trends in male birth proportion, data were classified according to maternal race/ethnicity (non-Hispanic white, non-Hispanic black, Hispanic); specific Hispanic ethnicity groups (e.g. Cuban, South American) were not analyzed.

Maternal age was categorized as <20, 20–24, 25–29, 30–34, 35–39 and ≥40 years. Although paternal age may have some independent association with sex ratio (Nicolich *et al*., 2000), US vital records are, on average, missing a significant proportion of records with paternal age (i.e. >10%) and paternal age varies greatly according to marital status depending on the state from which the birth records originate (Schoendorf and Branum, 2006). These limitations have reduced the usefulness of these administrative records for evaluating the role of paternal age on the US sex ratio.

For analyses by plurality, data were categorized as singleton or multiple (twins and higher) births. Gestational age was categorized as <28, 28–32, 33–36 and ≥ 37 weeks. Throughout the article the <28, 28–32 and 33–36 week groups will be collectively referred to as 'preterm' and births ≥37 weeks as 'term'. These categories were chosen so that differences in sex ratio among very preterm births (<28 weeks) could be separated from more moderately preterm, or term infants.

Statistical analysis

Trends in male birth proportion were assessed in two ways. First, we examined trends graphically using Stata (v. 10 SE) using LOWESS smoothing with a running-mean smoother and bandwith of 0.30. Crude trends were plotted alongside the smoothed trend lines in Fig. 1, in order to demonstrate the differences between these lines. Second, log binomial regressions were performed using individual year as the independent variable. Regression analyses were performed using PROC GENMOD in SAS (v. 9.1) using the 'link=log' function. Estimates from regression equations were plotted against actual male birth proportion from each year to check for goodness-of-fit in addition to using the *P*-values and confidence intervals to assess statistical significance. A value of $P < 0.05$ was considered significant. The beta coefficients derived from these regressions serve as the trend test i.e. the increase or decrease in male birth proportion per year over time. All analyses were conducted separately by race (white and black).

After assessing trends graphically and fitting log binomial models, an interaction term between year and the covariate of interest (i.e. plurality, gestational age or Hispanic ethnicity) was added to the log binomial model, where appropriate, to assess differences in trends among categories of the covariate. If no interaction was present, overall trends in male birth proportion were adjusted for the covariate of interest along with maternal age. If an interaction was present, male birth proportion trends were presented separately for each category of the covariate, adjusted for maternal age.

Results

Table 1 illustrates the proportion of births in the USA by plurality, maternal age, gestational age and Hispanic ethnicity for selected years

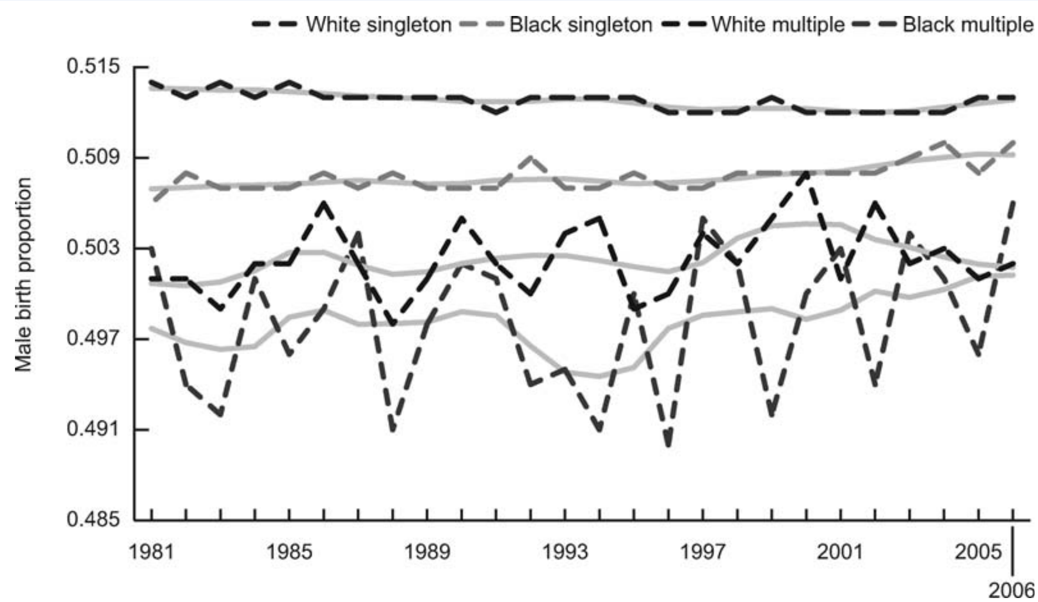

Figure 1 Trends in male birth proportion by plurality and race: 1981–2006.
Source: Centres for Disease Control and Prevention/National Center for Health Statistics (CDC/NCHS) US National Vital Natality Data: 1981–2006. *Smoothed line is represented by the solid line through the middle of each race and plurality specific crude trend.

among white and black births. Among white births, the percentage of multiple births, births to women 35 years of age and older, preterm births at 28–32 and 33–36 weeks, and births with reported Hispanic ethnicity increased between 1981 and 2006. Similarly, multiple births, births to women 30 years and older, births at <28 weeks, and births with reported Hispanic ethnicity increased among black births.

Trends in male birth proportion by year and differences according to the same characteristics are presented in Table II. Among white births the male birth proportion for all births included in the study decreased very slightly from 0.513 in 1981 to 0.512 in 2006. Among white births, the male birth proportion also decreased slightly, but not linearly, among singletons, women younger than 25 years and 35–39 years, term births and births at 28–32 weeks. Among white births in general, male birth proportion was lower among multiple births, compared with singletons, and higher for births in the preterm categories, compared with term births. White births with reported Hispanic ethnicity had a slightly lower male birth proportion compared with non-Hispanic ethnicity.

Among black births, male birth proportion increased non-linearly among all births, singletons, births to women 40+ years and among non-Hispanic births. Male birth proportion among black births was similar over categories of plurality and gestational age, although was higher among births at <28 weeks compared with births at other gestational ages and higher among births where mothers reported Hispanic ethnicity.

Has the increase in plural births affected the US sex ratio?

Smoothed and crude linear trends in male birth proportion by plurality among white and black births are illustrated in Fig. 1. As shown previously in Table II, there was a slight decrease in male birth proportion among white singletons only and non-linear increases among black singletons, and white and black multiple births (Fig. 1). Fitting log-binomial models to these data for white and black births with interaction terms to assess whether the trends differed by plurality, the interaction between year and plurality was significant for white births only ($P < 0.01$).

Due to the significant interaction, further analyses were separated by plurality for white births (Table III). Among white singletons, there was a significant decrease in male birth proportion over time ($P < 0.0001$). Adjustment for maternal age had no impact on these results, but when adjustment was made for maternal age and gestational age, the results indicate there would have been a greater decrease in male birth proportion over time (−0.0002 decrease in male birth proportion per year compared with −0.0001, $P < 0.0001$) had these two factors not changed. Among white multiples, there was an increase in male birth proportion over time, although this was not significant. When maternal age and gestational age were added to the model, the male birth proportion among white multiples decreased markedly ($P < 0.05$) and was similar of that among white singletons.

Among black births, adjustment for plurality had no impact on the significant increase in male birth proportion among all births. However, further adjustment for maternal age and gestational age resulted in a slightly greater increase in male birth proportion among black births over time (0.0002 increase in male birth proportion per year compared with 0.0001, $P < 0.0001$).

Have changes in the gestational age distribution affected the US sex ratio?

Among white births, the male birth proportion appeared to decline slightly over time, or remain stable for births at each category of

Table I Proportion of births in the USA by selected characteristics, year and race

	Year					
	1981	1986	1991	1996	2001	2006
White						
All *N*	2 919 869	2 994 233	3 212 499	3 081 601	3 166 455	3 296 759
Plurality						
Singleton	98.1	97.9	97.7	97.3	96.8	96.7
Multiple	1.9	2.1	2.3	2.7	3.2	3.3
Maternal age (years)						
<20	12.8	10.6	11.0	11.3	10.1	9.4
20–24	33.3	28.8	25.6	23.5	24.5	24.7
25–29	32.5	33.4	30.9	28.4	26.8	28.3
30–34	16.7	20.2	22.7	24.2	24.5	22.9
35–39	4.0	6.3	8.4	10.7	11.6	12.1
40+	0.6	0.8	1.3	2.0	2.5	2.7
Gestational age (weeks)						
<28	0.4	0.4	0.4	0.4	0.4	0.5
28–32	1.0	1.0	1.1	1.2	1.3	1.4
33–36	5.8	6.2	6.9	7.3	8.3	9.0
37+	88.5	88.5	86.8	85.7	84.6	84.1
Hispanic ethnicity						
Hispanic	[1]	[1]	18.8	22.1	26.1	29.6
Non-Hispanic	[1]	[1]	80.4	76.3	73.2	69.8
Black						
All *N*	555 631	583 204	675 940	590 347	602 539	662 556
Plurality						
Singleton	97.5	97.5	97.2	97.0	96.6	96.3
Multiple	2.5	2.5	2.8	3.0	3.4	3.7
Maternal age (years)						
<20	25.7	23.0	23.0	22.8	18.8	17.0
20–24	35.4	33.8	32.1	30.2	32.9	32.1
25–29	23.5	24.9	23.9	22.4	22.7	25.1
30–34	11.3	13.2	14.6	15.9	15.6	15.6
35–39	3.3	4.4	5.5	7.3	8.1	8.1
40+	0.7	0.7	0.9	1.4	1.9	2.1
Gestational age (weeks)						
<28	1.2	1.3	1.4	1.4	1.4	1.4
28–32	2.8	2.9	3.0	2.7	2.6	2.7
33–36	11.5	11.9	12.3	11.3	11.6	12.3
37+	79.9	79.3	76.1	77.2	77.1	76.6
Hispanic ethnicity						
Hispanic	[1]	[1]	1.4	2.0	2.3	6.7
Non-Hispanic	[1]	[1]	97.7	97.2	97.3	92.6

[1]Data on Hispanic ethnicity not available until 1989.
Source: Centres for Disease Control and Prevention/National Center for Health Statistics (CDC/NCHS) US National Vital Natality Data: 1981–2006.

gestational age (Fig. 2), although as stated previously, male birth proportion was higher among births in the preterm categories compared with term births. Although the interaction with year was significant for white births at each gestational age category except <28 weeks, indicating different trends in male birth proportion by gestational age, the actual estimates were generally in the same direction (trends among preterm births showed less of a decrease compared with term births). Therefore, results were not further stratified and

Table II Male birth proportion in the USA by selected characteristics, year and race

	Year					
	1981	1986	1991	1996	2001	2006
White						
All births	0.513	0.513	0.512	0.512	0.511	0.512
Plurality						
Singleton	0.514	0.513	0.512	0.512	0.512	0.513
Multiple	0.501	0.506	0.502	0.500	0.501	0.502
Maternal age (years)						
<20	0.516	0.514	0.511	0.514	0.512	0.512
20–24	0.514	0.512	0.512	0.512	0.511	0.512
25–29	0.513	0.513	0.511	0.512	0.512	0.513
30–34	0.512	0.513	0.512	0.512	0.511	0.513
35–39	0.512	0.512	0.510	0.513	0.510	0.510
40+	0.512	0.506	0.512	0.509	0.512	0.510
Gestational age (weeks)						
<28	0.546	0.549	0.541	0.543	0.538	0.538
28–32	0.546	0.540	0.540	0.537	0.538	0.532
33–36	0.541	0.538	0.537	0.543	0.536	0.537
37+	0.511	0.511	0.509	0.509	0.508	0.509
Hispanic ethnicity						
Hispanic	[1]	[1]	0.509	0.510	0.509	0.511
Non-Hispanic	[1]	[1]	0.512	0.513	0.512	0.513
Black						
All births	0.506	0.508	0.507	0.507	0.508	0.510
Plurality						
Singleton	0.506	0.508	0.507	0.507	0.508	0.510
Multiple	0.503	0.499	0.501	0.490	0.503	0.506
Maternal age (years)						
<20	0.508	0.512	0.511	0.508	0.509	0.511
20–24	0.506	0.508	0.507	0.507	0.507	0.511
25–29	0.506	0.508	0.505	0.507	0.510	0.510
30–34	0.504	0.506	0.507	0.504	0.507	0.507
35–39	0.502	0.501	0.506	0.506	0.503	0.507
40+	0.489	0.492	0.496	0.500	0.507	0.505
Gestational age (weeks)						
<28	0.522	0.535	0.531	0.530	0.534	0.526
28–32	0.510	0.507	0.506	0.504	0.512	0.507
33–36	0.512	0.511	0.512	0.512	0.516	0.519
37+	0.505	0.507	0.505	0.505	0.505	0.508
Hispanic ethnicity						
Hispanic	[1]	[1]	0.515	0.515	0.510	0.513
Non-Hispanic	[1]	[1]	0.507	0.506	0.508	0.509

[1]Data on Hispanic ethnicity not available until 1989.
Source: CDC/NCHS US National Vital Natality Data: 1981–2006.

we refer to the gestational age-adjusted results from the stratified plurality analysis for white births to illustrate the impact of adjustment for gestational age on overall trend in male birth proportion trends (Table III); as stated above, adjustment for gestational age further decreased male birth proportion among white singletons and reversed the trend for multiples.

Among black births, there were similar trends in male birth proportion according to gestational age category (Fig. 3). Interaction

Table III Log binomial regression coefficients (β_1) (95% confidence interval) modeling plurality and time trends in sex ratio in the USA (1981–2006)

	White singletons*		White multiples*	
	Intercept β_0	β_1 (Year)	Intercept β_0	β_1 (Year)
Crude	−0.6667	−0.0001[1] (−0.0001, −0.00007)	−0.6905	0.0002 (−0.00002, 0.0003)
Model 1[a]	−0.6667	−0.0001[1] (−0.0001, −0.00007)	−0.6890	0.00004 (−0.0001, 0.0002)
Model 2[b]	−0.6704	−0.0002[1] (−0.0002, −0.0001)	−0.7006	−0.0002[2] (−0.0004, −0.00002)
	Black births*			
Crude	−0.6801	0.0001[1] (0.0001, 0.0002)		
Model 1[c]	−0.6797	0.0001[1] (0.0001, 0.0002)		
Model 2[d]	−0.6809	0.0002[1] (0.0001, 0.0003)		
Model 2[e]	−0.6832	0.0002[1] (0.0001, 0.0003)		

*Includes both Hispanic and non-Hispanic ethnicities.
[a]Model adjusted for maternal age.
[b]Model adjusted maternal age and gestational age.
[c]Model adjusted for plurality.
[d]Model adjusted plurality and maternal age.
[e]Model adjusted plurality, maternal age, and gestational age.
[1]$P < 0.0001$ (refers to significant increasing or decreasing trend as estimated by β1).
[2]$P < 0.05$ (refers to significant increasing or decreasing trend as estimated by β1).

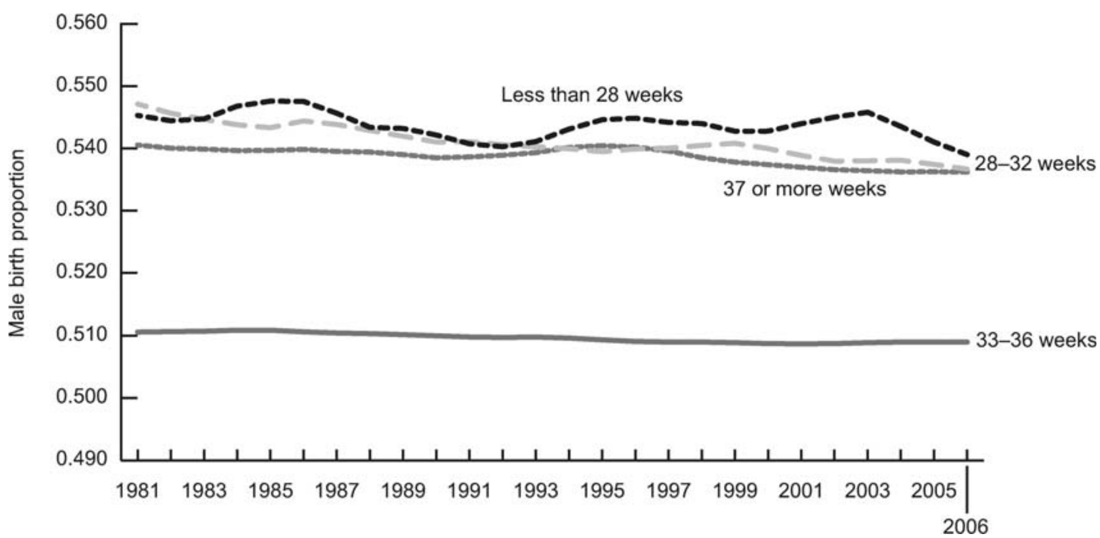

Figure 2 Trends in male birth proportion by gestational age among white births: 1981–2006.
Source: CDC/NCHS US National Vital Natality Data: 1981–2006.

between year and gestational age category was significant only among births at 33–36 weeks (data not shown). Similar to white births the trend at 33–36 weeks was in the same direction as term births, so the results were not further stratified and we again refer to the effect of adjustment for gestational age from Table III where adjustment for mother's age and gestational age increased the estimated trend.

Have increases in births to Hispanic mothers affected US sex ratio?

Since the time points for the Hispanic ethnicity analysis differed from the previous analyses, the interaction between year and plurality was reassessed for both black and white births. Given the shorter time frame, plurality no longer demonstrated a significant interaction with year among white births and was not significant for black births; therefore, the Hispanic ethnicity analysis was carried out among all white births and black births without stratification.

Among white births, after adjusting for Hispanic ethnicity, the decrease in male birth proportion was no longer significant (Table IV). However, after further adjusting for gestational age and maternal age, the decrease in male birth proportion among white singletons was significant and of the same magnitude as the crude model ($P < 0.001$). Among black births adjustment for Hispanic ethnicity had no effect on trends in male birth proportion over time. Further

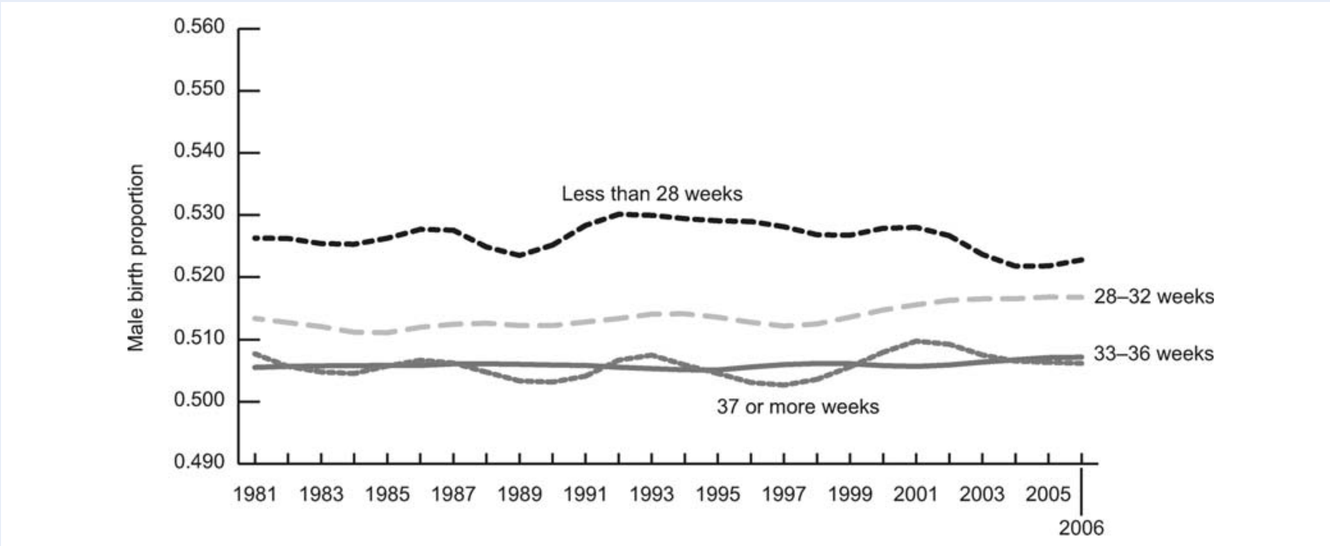

Figure 3 Trends in male birth proportion by gestational age among black births: 1981–2006.
Source: CDC/NCHS US National Vital Natality Data: 1981–2006.

Table IV Log binomial regression coefficients (95% confidence intervals) modeling Hispanic ethnicity and time trends in sex ratio in the USA (1989–2006)

	White		Black	
	intercept β_0	β_1	intercept β_0	β_1
Crude	−0.6684	−0.0001[1] (−0.0001, −0.0002)	−0.6797	0.0002[2] (0.00008, 0.0003)
Model 1[a]	−0.6772	−0.00004 (−0.0001, 0.0001)	−0.6757	0.0002[2] (0.0001, 0.0003)
Model 2[b]	−0.6776	−0.00003 (−0.0001, 0.0002)	−0.6760	0.0002[2] (0.0001, 0.0003)
Model 2[c]	−0.6823	−0.0001[2] (−0.0002, −0.0001)	−0.6774	0.0002[2] (0.0001, 0.0003)

[a]Model adjusted for Hispanic ethnicity.
[b]Model adjusted for Hispanic ethnicity and maternal age.
[c]Model adjusted for Hispanic ethnicity, maternal age, and gestational age.
[1]$P < 0.01$ (refers to significant increasing or decreasing trend as estimated by β1).
[2]$P < 0.001$ (refers to significant increasing or decreasing trend as estimated by β1).

adjustment for gestational age and maternal age also had no impact on trends (Table IV). Generally, results for this shortened time interval were consistent with those for the longer time interval for black births and for white singleton births (Fig. 4).

Discussion

Our analysis indicates inconsistent race-specific trends in the male birth proportion over time according to race, plurality, gestational age and Hispanic origin. Although none of these factors explained recent trends in the US male birth proportion, changes in the gestational age-specific birth distribution may have had some impact on tempering trends. The results also demonstrate that the recently described decrease in the overall US sex ratio is largely limited to the group comprising the largest number of births, white singleton infants born at term, and has leveled off in more recent years. This provides evidence that recent changes in the US sex ratio have occurred mainly among certain subpopulations, versus the general population.

One of the more interesting findings of this analysis concerns the differences in effects of changing gestational age and maternal age distributions on the trends in male birth proportion according to race. The trend of increasing male birth proportion associated with multiple births shown in the current analysis was unexpected given that there are some indications in the medical literature that twinning is associated with fewer male births, compared with singletons (James, 1975; Jacobsen et al., 1999), and that artificial ovarian stimulation may also lead to increased female births (Sampson et al., 1983; Dickey et al., 1995). However, differential changes in the gestational age structure according to both plurality and race may explain the divergent trends.

Between 1981 and 1998, births at 29–32 weeks increased 39% among white multiple births, compared with 7% among white

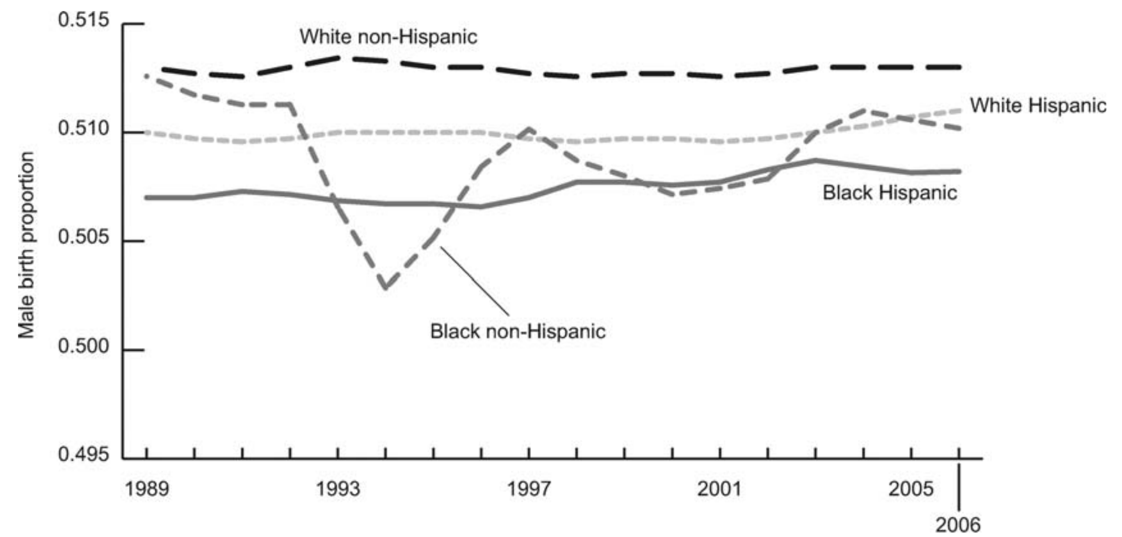

Figure 4 Trends in male birth proportion by race and ethnicity: 1989–2006.
Source: CDC/NCHS US National Vital Natality Data. Data on Hispanic ethnicity not available until 1989.

singletons (Branum and Schoendorf, 2002). In addition, births at 33–36 weeks increased 52% among white multiples compared with 29% among white singletons. As a result term births among white multiples decreased nearly 30% compared with only 2% among white singletons. Term multiple births have a lower sex ratio compared with preterm multiple births and term singleton births. Assuming the trends in preterm birth continued through 2006, the relatively large decrease in term births among white multiples resulted in an increase in male births. Adjustment for changes in the gestational age distribution explained the increase seen among white multiple births.

Changes in gestational age-specific births were similar for black multiples and white multiples, and births increased most among births occurring at 33–36 weeks whereas term births decreased 25% (Branum and Schoendorf, 2002). However, unlike white births, the sex ratio for black singleton and multiple births is very similar, which may explain why there was less of an effect of gestational age on trends in male birth proportion and no differences in trends according to plurality.

There were also differential effects of maternal age on male birth proportion according to race. Among white singletons, adjustment for maternal age had no additional impact on trends once gestational age was taken into account. However, among black births, maternal age had more impact on adjustment for trends in male birth proportion than gestational age. Male birth proportion tends to be slightly lower for older women among black births but not white (Table II). Therefore, the relatively large increases in births to older women, may have had more of an effect among black, compared with white, births.

There were also differential effects of Hispanic ethnicity on male birth proportion according to race. A previous analysis of sex ratio trends in California, demonstrated that Hispanic births had a lower sex ratio than non-Hispanic white births and that the increasing proportion of Hispanic births over time may have explained the apparent decrease in sex ratio among white births (Smith and Von Behren, 2005). For white births, this was demonstrated in the current analysis as well, although this effect was negated once gestational age was considered. Although the higher proportion of male births among black Hispanic births, compared with black non-Hispanic births, could have contributed to the increase in male birth proportion among black births, there is a smaller proportion of black births with Hispanic ethnicity compared with white births. However, Hispanic ethnicity has increased dramatically among black births in the last 6–7 years. Thus it will remain to be seen if Hispanic ethnicity has an impact on the sex ratio trend among black births in the future.

In addition to parental age and race, previous analyses of the sex ratio have demonstrated significant effects of a variety of factors, including, but not limited to, maternal smoking (Beratis *et al*., 2008), maternal diet (Rosenfeld and Roberts, 2004), socio-economic status (Almond and Edlund, 2007), geographic latitude (Grech *et al*., 2000), time to pregnancy (Smits *et al*., 2005) and environmental toxins (Vartiainen *et al*., 1999). Given the large populations in which these analyses are typically performed, it is not surprising that many significant relationships have been demonstrated between various factors and changes in sex ratio over time. This also illustrates the potential limitation of making individual-level inferences from population-level data in analyses carried out on a national or even regional level (Bonde and Wilcox, 2007). Future studies of sex ratio trends should be careful in making inferences about a single factor within a population, particularly those that are diverse with regards to race, geographic area or other demographic characteristics.

Although the current analysis demonstrates different effects of various factors on trends in sex ratio by race, it does not explain the apparent decrease in male birth proportion seen among white births and the increase among black births. The underlying causes of the race and ethnicity differences in the baseline sex ratio and in sex ratio trends, beyond those previously examined and discussed are uncertain and not likely to be explained by analysis of vital statistics data. Given the basis of 'race' and 'ethnicity' (David and Collins, 2007), especially as collected on birth certificates, it is probable that underlying differences in the sex ratio are more likely related to differential exposures rather than innate subgroup differences.

A limitation of this analysis, as with any analysis of gestational age data from vital statistics, is that there may be errors in the estimation of gestational age based on LMP. However, we did exclude implausible values prior to final analysis and, furthermore, the elimination of records based on a clinical estimate of gestational age will have limited possible bias. An important strength of this analysis is the use of comprehensive birth data over a 25 year period and the ability to measure Hispanic ethnicity on a national level, although less data were available for this analysis.

In summary, trends in the sex ratio at birth in the USA from 1981 to 2006 differ according to plurality, gestational age and race and ethnicity. The previously described decrease in the sex ratio among white births has stabilized in recent years and is largely driven by a decrease among white term singleton infants. Changes in the gestational age-specific birth distribution have had some impact on the sex ratio among white births, whereas changes in maternal age may have had more impact on the sex ratio among black births. The lack of a consistent pattern of change in the sex ratio across different subpopulations of US births suggests that a single mechanism is unlikely to explain the oft-referenced decrease in the overall US sex ratio. Instead, a variety of factors most likely contribute to the overall US trend. Elucidation of those factors, and the subsequent understanding of the mechanisms by which they act, will likely require smaller, more specialized cohorts, rather than large, generalized data sets, such as vital statistics, which have been the basis of this and most other US-based research.

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