Trends in the Probability of Twins and Males in California, 1983–2003

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This study examines the probability of twins by birth year, maternal race–ethnicity, age, and parity and the influences of these demographic factors on the probability of male in twins and singletons in a large, racially diverse population. Recent publications note steep increases in twin births while the probability of male births has been reported to vary by parental race–ethnicity and age and birth order. Probability of male stratified by plurality has not been investigated in California prior to this study. Cubic spline estimates and Poisson regression techniques were employed to describe trends in twins and males using California vital statistics birth and fetal death records over the period from 1983–2003. This study includes 127,787 twin pair and 11,025,106 singleton births. The probability of twins varied by birth year, maternal race–ethnicity, age, and parity. The probability of twins increased by 10.1% from 1983–1992 and increased by 20.1% from 1993–2003, nearly doubling the previous increase. All maternal race–ethnicity groups showed increases in probability of twins with increasing maternal age. Parous women compared to nulliparous women had larger increases in the probability of twins. The probability of males in twins decreased from 1983–1992 and increased from 1993–2003; while in singletons the probability appeared unchanged. These findings show increases in the probability of twins in California from 1983–2003 and identify maternal age, race–ethnicity, and parity groups most likely to conceive twins. The cause of the increase in twins is unknown but coincides with trends towards delayed childbearing and increased use of subfertility treatments.

Keywords: twins, maternal race–ethnicity, parity, age, probability of male

A substantial increase in twinning prevalence has been reported in the United States (Jones, 2003; Luke, 1994; Luke & Martin, 2004; Martin, 1999; Russell et al., 2003) and internationally (Chen et al., 1992; Eriksson & Fellman, 2007; Fellman & Eriksson, 2005; Eriksson & Fellman, 2004) over the past 2 decades. It is estimated that between one-fourth and one-third of the increase in multiple births in the United States is due to delayed childbearing until advanced maternal ages (Jewell & Yip, 1995; Luke & Martin, 2004; Lynch et al., 2001). Increases in twins have been reported with older maternal age (Fellman & Eriksson, 2005; Luke & Martin, 2004; Martin, 1999; Mosteller et al., 1981), among specific race–ethnicity groups (Luke, 1994; Luke & Martin, 2004; Martin, 1999; Mosteller et al., 1981; Pollard, 1995; Russell et al., 2003), and higher parity (Astolfi et al., 2003; Basso et al., 2004; Beemsterboer et al., 2006; Bonnelykke, 1990; Bulmer, 1970; Chen et al., 1992; Mosteller et al., 1981; Picard et al., 1991; Shipley et al., 1967).

The probability of male births, approximately 51.5% (Smith & Von, 2005) of births, has been observed to differ between twins and singletons (Beiguelman et al., 1995; James, 1975; Picard et al., 1991). Variations in the probability of males has been reported with parental age (James & Rostron, 1985; Nicolich et al., 2000), race–ethnicity (Nicolich et al., 2000; James, 1987), parity (James & Rostron, 1985a), and occupation (James, 2004; James & Rostron, 1985), stress levels (Davis et al., 2007), and specific subfertility treatments (Check et al., 1994).

Results from previous studies (Beiguelman et al., 1995; James, 1975; Picard et al., 1991) comparing the probability of males in twins and singletons are inconsistent, with some studies showing lower frequencies of males in twins as compared to singletons (James, 1975; Picard et al., 1991) while results from other studies (Beiguelman et al., 1995) have not supported this finding. These inconsistent findings may be in part due to limitations in statistical power due to small sample sizes. An investigation of the influences of demographic variables on the probability of males among twins and singletons, using a larger, more recent, and more diverse study population is needed. We therefore examine influences of selected demographic variables on trends in the frequency of twin births and trends in the probability of males among

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twins and singletons, using data from a racially and ethnically diverse study population of over 11 million births in California from 1983 to 2003.

Methods

Study Population

This study's population is twin and singleton subjects extracted from California vital statistics birth and fetal death records (fetal deaths at greater than 20 weeks gestation) for the years 1983–2003 (n = 255,574 twin and 11,025,106 singleton births). Births other than twins and singletons (n = 9,947 subjects) were excluded from analyses. Beginning with a dataset containing 263,507 twin subjects, linking criteria of mother's last name, mother's maiden name, and child's date of birth were used to match twin subjects and after removal of missing values, the dataset contained 262,551 twin subjects. When co-twin siblings could not be identified, the twin was eliminated, resulting in a loss of 6,583 twin subjects. Maternal identification for each twin born was impossible for 79 twin pairs with identical linkage criteria. The final dataset contained 127,787 twin pairs.

Explanatory variables examined were birth year, maternal race–ethnicity (White, Latina, African–American, and Asian), maternal age (categorized as 13–19, 20–24, 25–29, 30–34, and greater than or equal to 35 year olds), parity (categorized as none and one or more previous live births), plurality (twin and singleton), and sex of infant. All covariate information was obtained from California birth and fetal death certificates.

Statistical Methods

The probability of twins was defined as the number of twin subjects divided by the total of the number of twin and singleton subjects and included live and stillbirths. Probability of a male was calculated by dividing the numbers of males by the total of the numbers of twin and singleton subjects. In these analyses probability is used in preference to rate, because probability is a more appropriate measure for a phenomenon measured without reference to time at risk.

The trends in probability of twins were described by a least squares estimated line using a three quantile knot, restricted cubic spline curve within each maternal race–ethnicity, age, and parity grouping. Poisson regression techniques were used to assess the evidence of nonlinearity in trends of annual twin probabilities. Comparisons of the linear and nonlinear Poisson models were performed using likelihood ratio chi-square tests.

Statistical tests were not employed where the influence of sampling variation was essentially negligible due to the extremely large number of singleton and twin births in California during the years 1983–2003. However, statistical tests and models were employed only when the data were stratified by more than three variables where the sample size was substantially reduced. All figures presented illustrate statistically significant model results (two-sided p value < .01) when Poisson regression was employed.

Results

Twin probabilities

The frequency of twin and singleton births vary by maternal age and race–ethnicity (Table 1). Twins in Whites (48.8%) represent approximately half of all twins born in California from 1983 to 2003. The probability of twins in African–Americans (0.0293) and Whites (0.0269) were higher compared to Latinas (0.0183) and Asians (0.0177). The maternal age distributions vary between race–ethnicity groups with 55% and 63% (respectively) of twins born to mothers over age 29 in Whites and Asians; while in African–Americans and Latinas 33% and 35% (respectively) of twins are born to mothers over age 29.

Twin probabilities range from 0.019 in 1983 to 0.028 in 2003 in California from 1983–2003. A piecewise single knot spline estimated line (Figure 1), indicates that the twin probability increased by 10.1% from 1983 to 1992, while the twin probability increased by 20.1% from 1993–2003.

The annual probability of twins by maternal age and race–ethnicity groups is displayed in Figure 2 and in Whites increases with birth year and maternal age, with the greatest increase among mothers 35 years of age and older. The probability of twins in Latinas increases modestly with maternal age in comparison to other race–ethnicity groups. Probabilities of twins among African–Americans increase with year of birth and in all age groups. The greatest change is observed among the oldest mothers. Generally the distribution of twin probabilities in Asian mothers is similar to the one observed in Whites, although lower in every maternal age group.

The largest per cent change in twinning probabilities were in Asians and Whites 35 years of age and older (both increased 157.4%), while decreases over the same period were observed among Asian women less than 24 years of age (Table 2).

Comparison of parous and nulliparous groups show that White and Asian parous women 35 years or older have the largest increases in the likelihood of a twin birth while nulliparous women 34 years of age or younger tended to have slight decreases in the change in the probability of a twin birth over time (Figure 3). Among nulliparous women, the likelihood of twin births during the 21-year period is similar across the four maternal race–ethnicity groups. Among parous women distinct differences in the probability of twins are evident among maternal race–ethnicity groups. The greatest change in the likelihood of a twin birth was observed in White women 35 years of age or older. Generally, the trends in the likelihood of twin births when the data are stratified by maternal race–ethnicity, age, and parity groups are similar to those observed in Figure 2. The twin probabilities in the parous groups, regardless of maternal race–ethnicity, are approximately twofold higher compared to nulliparous women and generally increase from 1983 to 2003.

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Table 1

Frequency of Twin and Singleton Births by Maternal Age and Race Categories, California (1983–2003)[a]

	Maternal age (years)						
	13–19	20–24	25–29	30–34	35–55	Missing	Total
Twin pairs							
Maternal race–ethnicity							
Whites	2,174 (3.7)	8,731 (14.8)	15,730 (26.6)	18,193 (30.8)	14,229 (24.1)	23 (0.04)	59,080 (48.8)
African–Americans	1,450 (11.2)	3,520 (27.3)	3,655 (28.3)	2,796 (21.7)	1,471 (11.4)	8 (0.06)	12,900 (10.7)
Latinas	4,257 (9.9)	11,207 (26.1)	12,390 (28.9)	9,456 (22.0)	5,573 (13.0)	13 (0.03)	42,896 (35.4)
Asians	162 (2.6)	643 (10.2)	1,515 (24.1)	2,219 (35.3)	1,744 (27.8)	2 (0.03)	6,285 (5.2)
Total	8043 (6.6)	24,101 (19.9)	33,290 (27.5)	32,664 (27.0)	23,107 (19.1)	46 (0.04)	121,161 (100.00)
Singletons							
Maternal race–ethnicity							
Whites	318,694 (7.5)	928,389 (21.7)	1,272,589 (29.8)	1,125,815 (26.3)	630,514 (14.7)	1,053 (0.03)	4,277,054 (41.0)
African–Americans	148,913 (17.4)	249,419 (29.2)	214,102 (25.1)	141,460 (16.6)	72,714 (8.5)	690 (0.08)	854,298 (8.2)
Latinas	706,654 (15.3)	1,416,862 (30.7)	1,274,278 (27.6)	797,076 (17.3)	414,508 (9.0)	1,908 (0.04)	4,611,286 (44.2)
Asians	32,257 (4.6)	92,078 (13.2)	204,351 (29.3)	229,978 (33.0)	137,512 (19.7)	338 (0.05)	696,514 (6.7)
Total	1,206,518 (11.6)	2,686,748 (25.7)	2,965,320 (28.4)	2,294,329 (22.0)	1,255,248 (12.0)	3,989 (0.04)	10,439,152 (100.0)

Note: [a]Figures in parentheses are percentages out of corresponding row or column total; for example, the total percentage of White twins is as follows: [(total number of White twins)]/[(total number of twins)] × 100 = [(59,080)]/[(121,161)] × 100 = 48.8.

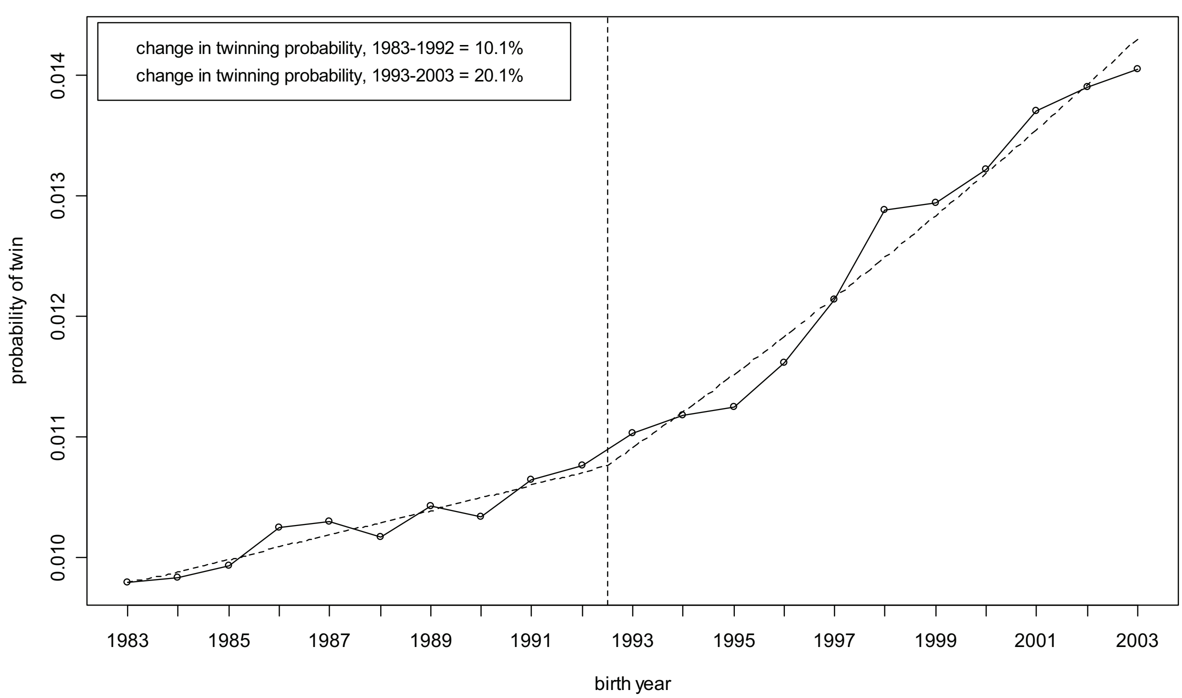

Figure 1

Probability[a] of twins by birth year, California (1983–2003).

Note: [a]Probabilities calculated as follows: [(number of twin subjects born each year) / (number of twin and singleton subjects born each year)]. Horizontal dashed line represents one knot linear piecewise spline. Solid circles represent observed data.

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Table 2

Per Cent Change in Twinning Probability for Five Maternal Age and Four Maternal Race–Ethnicity Groups, California (1983–2003)

	[a]Per cent change in twinning probability (%)			
Age group (years)	Whites	African–Americans	Latinas	Asians
13–19	24.2	41.0	12.9	–29.7
20–24	23.1	45.2	7.1	–18.6
25–29	47.3	66.5	11.3	28.0
30–34	77.0	31.5	10.1	61.7
35–55	157.4	38.0	10.6	157.4

Note: [a] For each race–ethnicity and age group per cent change calculated as follows: [(probability of twins in 2003)/(probability of twins in 2003 — probability of twins in 1983)] * 100

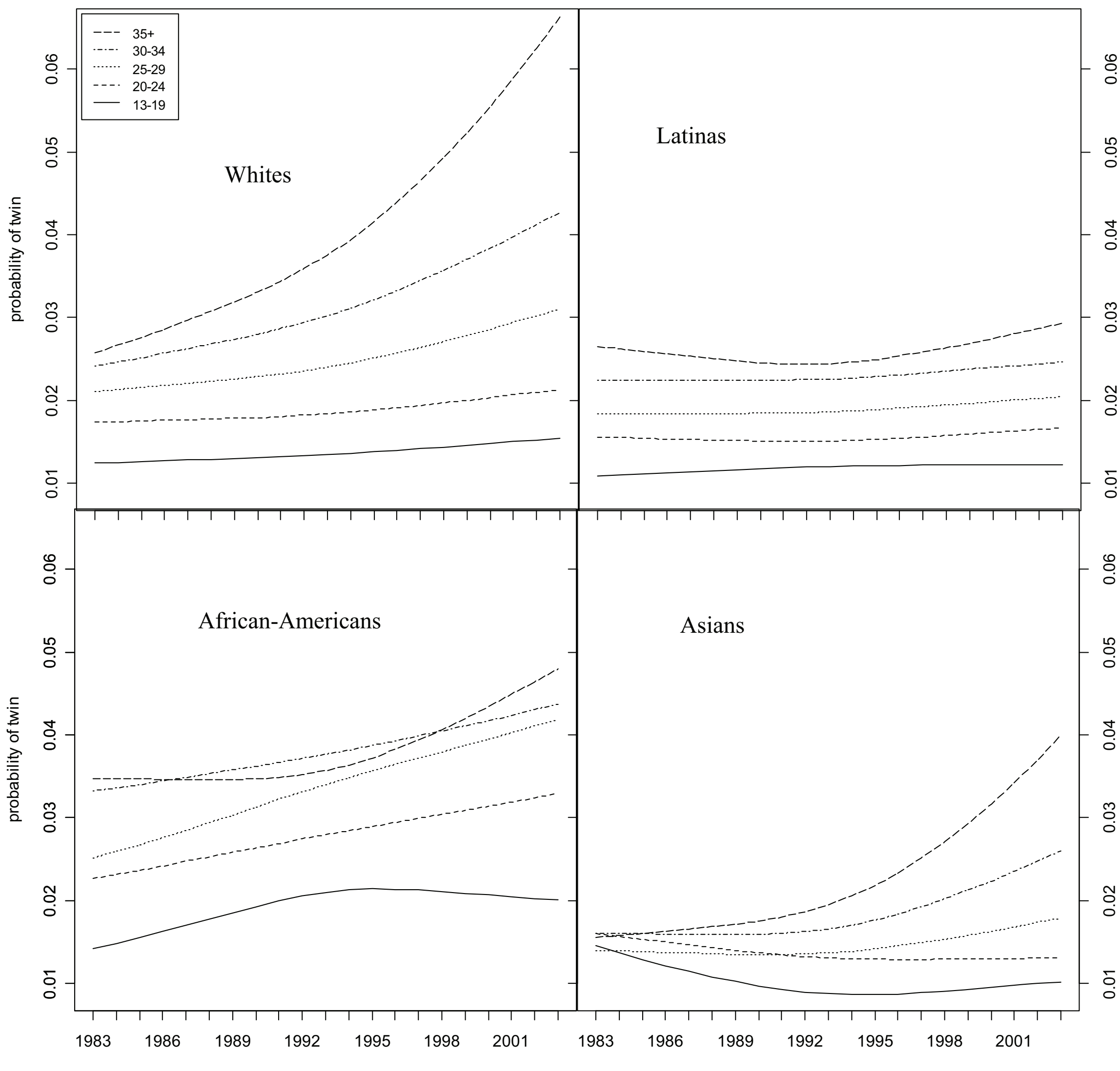

Figure 2

Probabilities[a] of twins stratified by maternal age and race–ethnicity, California (1983–2003).

Note: [a]Probabilities calculated as follows: [(twin subjects born each year in each maternal race–ethnicity/age group) / (twin subjects and singleton subjects born each year in each maternal race–ethnicity/age group)]. Curves estimated with restricted cubic splines.

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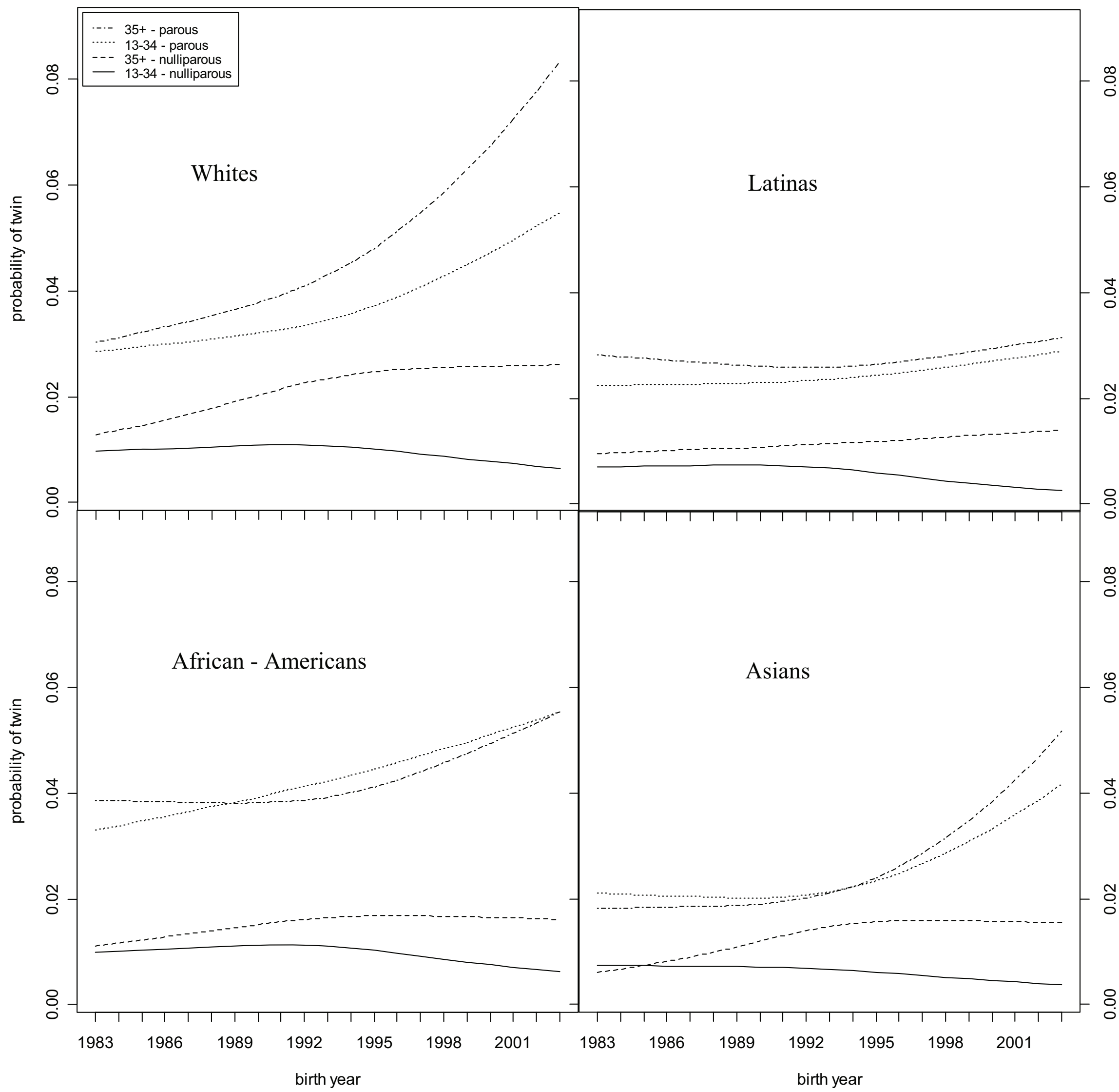

Figure 3

Probabilities[a] of twins stratified by maternal age, race–ethnicity, and parity, California (1983–2003).

Note: [a]Probabilities calculated as follows: [(twin subjects born each year in each maternal parity group) / (twin subjects and singleton subjects born each year in each maternal parity group)]. Curves estimated with restricted cubic splines.

Probabilities of Males

The trends in the probability of a male in twins and singletons over the 21-year time period differ, with substantially more variability observed in twins. The probability of a male among twins appears to decrease (Figure 4) from 1983 to a minimum in 1992 and then increase until 2003; while the probability of a male among singletons appears stable and higher during this time period. The probability of males among twins was lower among mothers who were 13–34 years old (0.499) than among mothers greater than 35 years of age (0.511) regardless of parity status, while the probability of males among singletons was almost identical for all maternal ages and parity status (0.511) (detailed data not shown).

Two multivariate Poisson regression analyses describe the joint influences of birth year, maternal age, race–ethnicity, and parity on the probability of males among twins (Table 3) and among singletons (Table 4). A comparison of nonadditive and additive Poisson models indicated that the simpler additive models were adequate representations of the twin and singleton data. That is interaction effects were tested and found not statistically significant.

Comparison of birth year as linear and nonlinear influence (spline) on the probability of male twins suggests an improvement in model fit (two-sided p value = .07) using birth year as a nonlinear term. Significant differences were observed among maternal race–ethnicity groups (two-sided p values ≤ .05); while maternal age and parity had no substantial influence on the

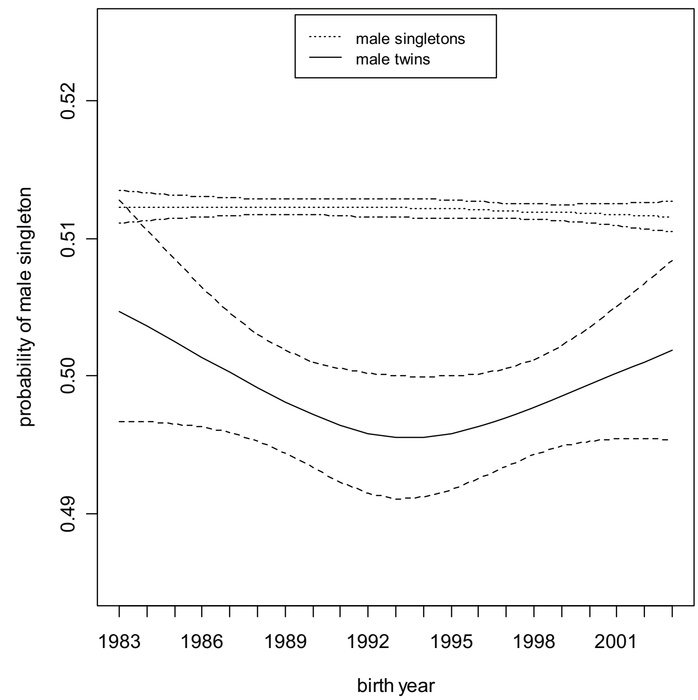

Figure 4

Probability of male twins[a] and singletons[b] with 95% confidence intervals (dashed lines) by birth year in California (1983–2003).

Note: [a] Probabilities calculated as follows: [(male twin subjects born each year) / (twin subjects born each year)].
[b] Probabilities calculated as follows: [(male singleton subjects born each year) / (singleton subjects born each year)].

probability of males among twins. Latinas and African–Americans have significantly reduced (0.97 and 0.98, respectively) probability ratios compared to Whites, while Asians who have a significantly increased (1.03) probability ratio adjusted for maternal age, parity, and year of birth.

Significant influences were observed for maternal race–ethnicity groups and parity (two-sided *p* values ≤ .05) for males among singleton births. Maternal age and birth year appeared not to influence the probability of males among singleton births. A significantly reduced (0.996) ratio of the probability of a male among singletons was observed for nulliparous women as compared to parous women after adjustment for birth year, maternal age, and race–ethnicity. Latinas and African–Americans illustrate significantly lower (0.994 and 0.990 respectively) ratios of probabilities of male singleton births compared to Whites,

Table 3

Poisson Results for Model[a] of Probability of Male Twin

	Estimate	Ratio[b]	95% confidence interval	*P* value
Intercept	4.1217	—	—	—
Birth year, 1st knot[c]	–0.0024	—	—	.08
Birth year, 2nd knot[c]	0.0029	—	—	.07
Age[d]	0.0101	1.01	1.00, 1.03	.14
Parity[e]	–0.0009	1.00	0.98, 1.02	.92
Latina[f]	–0.0267	0.97	0.96, 0.99	< .001
African–American[f]	–0.0190	0.98	0.96, 1.00	.05
Asian[f]	0.0301	1.03	1.01, 1.06	.02

Note: [a]Model is: log(probability of male|twin) = intercept + β 1 (cubic spline[birth year]) + β 2 (maternal age) + β 3 (maternal parity status) + β 4 (maternal race–ethnicity).
Chi square *p* value = .99 for comparison of interaction (all pairwise) and additive models. Chi square *p* value = .07 for comparison of additive model using birth year as linear and nonlinear influences.
[b]Rate ratios reflect per cent change.
[c]Birth year is represented as a nonlinear (spline) term.
[d]13–34 year old mothers as baseline.
[e]Parous mothers as baseline.
[f]White mothers as baseline.

Table 4

Poisson Results for Model[a] of Probability of Male Singleton

	Estimate	Ratio[b]	95% confidence interval	*P* value
Intercept	–0.6809	—	—	—
Birth year	–0.000027	1.000	0.996, 1.004	.72
Age[c]	–0.0007	0.999	0.997, 1.002	.60
Parity[d]	–0.0046	0.996	0.994, 0.997	< .001
Latina[e]	–0.0065	0.994	0.992, 0.995	< .001
African–American[e]	–0.0101	0.990	0.987, 0.993	< .001
Asian[e]	0.0047	1.005	1.001, 1.008	.01

Note: [a]Model is: log(probability of malel singleton) = intercept + β 1 (birth year) + β 2 (maternal age) + β 3 (maternal parity status) + β 4 (maternal race–ethnicity). Chi square *p* value = .22 for comparison of interaction (all pairwise) and additive models. Chi square *p* value = .50 for comparison of additive model using birth year as linear and nonlinear influences.
[b]Rate ratios reflect per cent change.
[c]13–34 year old mothers as baseline.
[d]Parous mothers as baseline.
[e]White mothers as baseline.

while Asians show a significantly increased (1.003) ratio of the probability of a male singleton birth compared to Whites.

Discussion

The probability of twins increased by 10.1% from 1983 to 1992 and increased by 20.1% from 1993 to 2003, nearly doubling the previous increase. A 52.9% increase in twin birth ratios was reported (Russell et al., 2003) using US live birth data from 1980 to 1999. Luke (Luke & Martin, 2004) observed a 77% increase in the number of twin births from 1980 to 2001. A 43.5% increase in the percentage of twin births from 1993 to 2000 was illustrated using data from Virginia (Jones, 2003). These accumulated findings indicate nationwide increases in twin births.

Our study includes data from live birth and fetal death certificates, while other studies (Jones, 2003; Luke, 1994; Martin, 1999; Russell et al., 2003) excluded fetal death data. Strengths of this study include use of a large dataset; in 2003 California represented approximately 12.2% of the US population (2006). California is a racially and ethnically diverse state, which allows for analysis of Latinas and Asians not available in other datasets that have reported solely on differences in African–Americans and Whites (Luke, 1994; Russell et al., 2003).

Limitations of our study include the inability to capture fetal deaths earlier than 20 weeks gestation. Our study lacks subfertility treatment use information, which could influence the probability of twins. Therefore our data in combination with other data allows for speculations but no conclusive evidence regarding the influence of subfertility treatments.

Our analysis includes the time period when subfertility treatments became widely used, specifically in vitro fertilization (Reynolds et al., 2003; Wright et al., 2007). The prevalence of use and types of subfertility treatments and procedures used to become pregnant were evaluated by interviewing mothers of male, liveborn nonmalformed subjects (n = 1,286) in The National Birth Defects Prevention Study (NBDPS) (Yoon et al., 2001). The NBDPS, a population based, case-control study, found that 3.8% of controls used subfertility treatments and procedures. Using this study's estimates (Yoon et al., 2001; Carmichael et al., 2007) we can expect 15,224 (11.9%) twin pairs and 410,061 (3.7%) singleton subjects due to these treatments in our data. While subfertility treatments and procedures may underlie some of the increase in twin births, it is unlikely that it completely explains the observed increase. The collection of expanded health data from the modified certificate of live birth (Martin & Menacker, 2007) will allow understanding of the influences of subfertility treatment practices on the demographics of the US population.

Stratification of our data by parity, maternal age, and race–ethnicity groups shows higher and greater increases in twin probability among parous women regardless of maternal race–ethnicity and age. A study by Astolfi (Astolfi et al., 2003), using data from Italy during 1951–1952, 1979–1981, and 1994–1996, showed three distinct trends in twin births by parity, maternal age, and time period. In 1951–1952, the analysis (Astolfi et al., 2003) showed increases in twin births with increasing parity reaching a maximum between 35–40 years of maternal age. Bulmer (Bulmer, 1970) also found twinning increased with increased parity reaching a maximum between 35–40 years of maternal age, which agrees with our data. Astolfi's analysis (Astolfi et al., 2003) of the 1994–1996 period illustrated rising twin births in nulliparous women that reached a maximum between 35–40 years of maternal age, while the numbers of twins born to third and higher birth orders, were constant in women less than 35 years of age and then decreased among women 35 years of age and older. While the 1994–1996 results do not agree with our findings the discrepancy may be attributable to increased use of

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subfertility treatments in Italy as we illustrate above that it is unlikely that many births in California can be credited to the use of subfertility treatments. Using Danish data from 1998–2001 the authors (Basso et al., 2004) noted that the percentage of twin deliveries was greater among multiparous compared to primiparous women of all ages for time to pregnancy of 12 months or less. A study (Bonnelykke, 1990) that used data from the Danish Medical Birth Register from 1984–1985 reported increases in dizygotic and decreases in monozygotic twins with increasing parity adjusted for maternal age. A study (Mosteller et al., 1981) using data from Virginia from 1960–1974, showed among Whites, increasing maternal age and birth order resulted in increased dizygotic twinning rates. A recent study (Beemsterboer et al., 2006) found older women (mean age 36.1 years) have a higher prevalence of multifollicular development due to the pituitary releasing FSH in response to decreased negative feedback induced by impending ovarian failure. This results in development of multiple follicles, and in the presence of two oocytes a dizygotic twin pregnancy results. Although our analyses did not examine zygosity, we observe increases in twinning in older maternal age groups, which could suggest that delayed childbearing until older maternal ages may underlie some of the observed increase in twinning.

Increased twinning could occur among nulliparous older mothers due to increased usage of subfertility treatments, however our results do not support this. Our data show the largest increases in the probability of twins in parous women 35 years of age and older and it seems unlikely that older parous women would comprise the majority of subfertility treatment users.

In our analyses similar trends are observed in the increases of the probability of twins by maternal age for three of the race–ethnicity groups, while we do not observe a similar trend in Latinas. The lack of increase in twin births among Latina women is an interesting finding and is in agreement with a study (Luke & Martin, 2004) that found fewer numbers of twins among all Hispanic maternal age groups using US data from 1996 to 2000.

Our findings regarding greater variation in the probability of males in twins than singletons agree with a Brazilian study (Beiguelman et al., 1995) examining twin births from 1984–1993. Our results are similar to other studies findings (James, 1975; Picard et al., 1991) , that reported lower probabilities of males in twins as compared to probabilities of males in singletons. Two studies (Beiguelman et al., 1995; Picard et al., 1991) reported lower probabilities of males in twins with increasing maternal age, while the study using data from several countries (James, 1975), showed that the sex ratio in twins appeared to increase until maternal age 35 and then decrease with mothers aged 40 to 44 years. In this analysis lower probability of male in twins was observed with mothers 35 years of age and older (data not shown). Higher parity was associated with increased probability of a male twin in the Israeli study (Picard et al., 1991) but was confined to the Jewish population. Bulmer (Bulmer, 1970) suggested that multiple pregnancies have high rates of fetal death and that fetal deaths have high sex ratios, which in turn lowers the sex ratio in liveborn twins.

A recent publication (Smith & Von Behren, 2005) regarding trends in the sex ratio of California births regardless of plurality, reported an overall male birth proportion between 0.510 and 0.514 per year from 1960 to 1996, with an average of 0.5122, which the authors compared to the average value of 0.5123 for the entire United States from 1958 to 1997. The authors reported decreases in the male proportion with increasing parental ages, birth order, and African–American and Native American race–ethnicity and conversely, increases in the male birth fraction with Asian race–ethnicity. The authors conclude the fluctuations in the male birth fraction are attributable to changes in the demographics of California.

Our findings concerning the probability of a male in singletons and twins agree with Smith and Von Behren's results (Smith & Von Behren, 2005). In our analyses statistically significant results in the probability of males in singletons and twins were associated with a reduced probability of a male in Latinas and African–Americans and an increased probability among Asians. Nulliparity was associated with a statistically significant reduced risk of probability of a male among singletons but not twins. Why parity adjusted for maternal age would be associated with probabilities of males in singletons but not in twins remains an unanswered question.

In conclusion, this study illustrates an increase in the probability of twins of 10.1% from 1983 to 1992 and an increase of 20.1% from 1993 to 2003, nearly a doubling of the previous increase. All maternal race–ethnicity groups showed increases in the probability of twins with increasing maternal age. Parous women compared to nulliparous women had larger increases in the probability of twins. The probability of males in twins decreased from 1983 to 1992 and increased from 1993 to 2003; while in singletons the probability appeared unchanged.

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