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# Sibling Deidentification

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A commonplace phenomenon, judgments of being different from one's sibling, was called *sibling deidentification*. We studied its possible determinants and attempted to clarify what subjects mean when they judge siblings to be different. Results on alike–different judgments elicited from 383 undergraduates, 203 from three-child and 180 from two-child families, show significantly higher levels of deidentification for *first pairs*, firstborn subject judging second-born sibling (S1Sb2) or second-born subject judging first sibling (S2Sb1), than for *jump pairs* (S1Sb3 or S3Sb1); with intermediate levels for *second pairs* (S2Sb3 or S3Sb2). There were no significant within-pair effects of birth order, sex of subject, and spacing. For three-child first pairs, same-sex siblings deidentify significantly more than opposite-sex siblings do. Results on semantic differentials indicate that when subjects judge their sibling different they mean that they polarize on significantly more personality dimensions than do those judging alike. Findings in conjunction with identification theory suggest that deidentification may be a mechanism for resolving sibling rivalry, a Cain Complex.

"My sister is entirely different from me." "My two children are as different as night and day." The present study of judgments of being different from one's sibling, what we call *sibling deidentification*, grew out of the strikingly high incidence of these remarks in informal everyday observation. Acquaintances, when asked what their siblings were like, or parents, when asked what their "other child" was like, very often produced statements of this kind. The phenomenon seemed particularly interesting in view of the seeming parallel with child–parent identification. Statements describing siblings as "entirely different" seemed to be as common as those describing children as being "just like" their parents. It was this parallel with identification that suggested the term *sibling deidentification*.

The use of the term deidentification need not imply that its determinants are the same as those for identification, especially since there are no less than three conflicting theories concerning the determinants of identification: psychoanalytic, social learning (Bandura, 1962), and cognitive–developmental (Kohlberg, 1966). While these theories could serve as fertile sources of hypotheses concerning determinants of sibling deidentification, there could be other sources as well. Specifically, the sociological studies of Bossard and Boll (1956) on personality role differentiation in large families (more than six siblings) seemed worth considering. Bossard and Boll attributed this role differentiation to the child's insistent striving for some special mark of distinction vis-à-vis his or her many siblings; the larger the family, the greater the likelihood of role differentiation. We planned to consider these alternative theories in connection with our data, but, until we know more about sibling deidentification, we intend to use the term in its most general sense, to refer to judgments of siblings being different from oneself.

Two issues were examined in the present

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study. First, we aimed to shed light on possible determinants of sibling deidentification in two- and three-child families. To this end, we examined variations in the percentage of deidentification judgments between and within sibling pairs. Between-pair comparisons in percent deidentification were undertaken in relation to three kinds of sibling pairs: *first pairs*, firstborn subject judging second sibling (S1Sb2) or second-born subject judging first sibling (S2Sb1); *second pairs*, second-born subject judging third sibling (S2Sb3) or third-born subject judging second sibling (S3Sb2); and *jump pairs*, firstborn subject judging third sibling (S1Sb3) or third-born subject judging first sibling (S3Sb1), the word *jump* alluding to the bypassed middle sibling. Within-pair comparisons in percent deidentification were undertaken in relation to the following pair characteristics: birth order; sex of subject; sex of sibling, male vs. female and same sex versus opposite sex; and spacing. This inclusive list of sibling variables was selected to provide a rich base for generating hypotheses.

Second, we attempted to gain some insight into what subjects mean when they judge their siblings to be different from themselves. To do so, we compared scores on semantic differential scales with global judgments of being different.

## Method

### Subjects

The total sample consisted of 383 students at Columbia University, 203 from three-child families and 180 from two-child families, the former providing deidentification judgments on first, second, and jump pairs, the latter providing an additional sample of first-pair judgments to be compared with the second and jump pairs of the three-child family. Students in large undergraduate classes at Barnard and Columbia College were recruited through their professors who were asked to give 10 minutes of their class time, usually at the end of class, for soliciting volunteer subjects and collecting data. A questionnaire was distributed to those volunteers who met the following criteria: member of intact family; both parents living; and neither parent previously married. The criteria were intended to eliminate half-siblings and to insure that families be intact. Twins were also excluded, as were siblings of subjects already included in the sample. Large undergraduate classes were solicited to facilitate data collection. The third-born subjects, who are generally underrepresented at college (Altus, 1966), were additionally recruited from door to door in the college dormitories. Table 1 shows the number of judgments of siblings in relation to each of the sample characteristics. It can be seen that the total number of judgments was large enough for between- and within-pair comparisons of percent deidentification.

Identifying data were solicited to ensure that all four pairs were matched in social class, ethnicity, and years at the university; 339 subjects responded. The total sample consisted of 54.6% professional, managerial, or technical; 41.2% skilled laborer; and 4.2% unskilled laborer. Ethnic composition was 53.4% Jewish, 25.7% Protestant ethnic (Scandinavian, Scottish, etc.), 11.3%

**Table 1: Number of Judgments for Pairs and for Members within Pairs in Relation to Sample Characteristics**

| | 3-child pair | | | | | | 2-child pair | |
|---|---|---|---|---|---|---|---|---|
| | First | | Second | | Jump | | First | |
| Characteristic | Prior (S1Sb2) | Later (S2Sb1) | Prior (S2Sb3) | Later (S3Sb2) | Prior (S1Sb3) | Later (S3Sb1) | Prior (S1Sb2) | Later (S2Sb1) |
| Male subject | | | | | | | | |
| Male sibling | 19 | 15 | 17 | 10 | 8 | 9 | 22 | 15 |
| Female sibling | 15 | 15 | 12 | 12 | 26 | 13 | 24 | 19 |
| Female subject | | | | | | | | |
| Male sibling | 28 | 18 | 18 | 17 | 37 | 17 | 28 | 19 |
| Female sibling | 28 | 16 | 16 | 10 | 19 | 10 | 26 | 27 |
| Spacing (months) | | | | | | | | |
| *M* | 32.6 | 41.2 | 53.6 | 53.2 | 84.3 | 86.9 | 41.4 | 48.7 |
| *SD* | 17.8 | 21.9 | 27.6 | 35.4 | 34.6 | 40.5 | 21.2 | 24.0 |
| *Mdn* | 29 | 35 | 49 | 44 | 82 | 81 | 36 | 44 |
| Pair totals | 154 | | 112[a] | | 139 | | 180 | |

*Note.* $N$ = 383. S1, S2, S3 = first-, second-, and third-born subjects, respectively; Sb1, Sb2, Sb3 = first-, second-, and third-born siblings, respectively.

[a] Missing one second born's judgment of third-born sibling.

Catholic ethnic (Italian, Irish, Polish, etc.), and 9.6% Minority ethnic (Black, Puerto Rican, Oriental, etc.) Regarding years at the university, 5.1% of the total sample were graduate students or enrolled at the Columbia School of General Studies, so that they were somewhat older than undergraduates. The remaining 94.9% were undergraduates, mainly Barnard females and Columbia males. Mean years at college for the undergraduates was 2.3. Tests for differences among the four pairs, chi-squares for social class and ethnicity and analysis of variance for years at the university, showed no significant differences, indicating adequate matching for these variables.

## Procedure

The questionnaire consisted of four parts, the first requesting birth date, sex, and age of the subject and each of his or her siblings; the second asking the subject to judge for each sibling, "Are you alike or different from your sibling?", and also to judge which parent[1] the subject and his or her sibling(s) resembled; the third consisting of a set of semantic differentials for each member of the family including the subject; and a final part for identifying data. The order of presentation of the second and third parts as well as the order of the sibling and parent questions of the second part were randomized across subjects.

**Global judgments.** The answer "different" to the question "Are you alike or different from your sibling?" was adopted as our measure of deidentification. In contrast to difference scores based on semantic differential scales, the global judgment "different" provides a measure of deidentification with optimal face validity, in view of the fact that deidentification is defined as judging the sibling as different. For both global judgments and semantic scales, subjects were instructed to respond in terms of personality characteristics, to prevent responding on the basis of physical characteristics.

Reliability data on the alike–difference question were available for stability of judgment over time. A sample of 72 students from two-child families was asked the question the previous semester in a preliminary pilot study. Eighteen of these subjects volunteered again in the present sample. Test–retest reliability was 83.3% agreement for the alike–difference question under study, indicating adequate stability for the measure of deidentification. In addition, effects of order of presentation on the answer to the question were studied by means of chi-square tests; none of the pairs showed significant order effects.

**Semantic differential.** On the basis of the pilot study, the semantic differential was adopted to study what subjects might mean when they judge siblings to be different. We had asked the pilot subjects to list three characteristics where they were alike or different from their sibling. For subjects with global judgments of different, 75% of the answers took the form of bipolar adjectives of the kind found on semantic differential scales (e.g., "He's an optimist, I'm a pessimist"). Further, the contents of all but a few of these characteristics seemed to be adequately covered by the standard sets of 7-point bipolar semantic differential items described in Osgood, Suci, and Tannenbaum (1957, chap. 2) and Snider and Osgood (1969, chap. 2). Relying on standard sets, especially that of Lazowick (1955) designed for an identification study, we selected the following items: happy–sad, pleasant–unpleasant, and good–bad (Osgood's evaluation factor); active–passive, fast–slow, and hot–cold (activity factor); strong–weak, rugged–delicate, and deep–shallow (potency factor); and tense–relaxed. To this standard set, we added those bipolar dimensions listed by our pilot subjects but not covered adequately by this set, namely, introvert–extrovert, conventional–unconventional, and achieving–nonachieving, making a total of 13 items. Items were presented in random order, with different randomizations for each member of the family to be rated.

To gain some insight into the meaning of judgments of being different, we planned to study these bipolar items in two stages. First, we focused on the total number of items where the subject polarized in relation to his or her sibling, where self and sibling ratings fell on opposite sides of the neutral midpoint for the item (e.g., self scored extrovert; sibling scored introvert). The total of these polarized items was called the Polarization (P) score. If the P score were found to be significantly associated with the global alike–difference judgment, we planned to proceed with an item analysis to see which of the items discriminated between subjects judging siblings as different and those judging alike. In addition, since some of these items were clearly sex-stereotyped (e.g., active–passive) while others were not (e.g., introvert–extrovert), we felt it was essential to evaluate whether polarization on sex-stereotyped dimensions contributed to the global judgment of different for opposite-sex siblings any more than it did for same-sex siblings. Item polarization was also examined in relation to differences in sex of subject.

Altogether, there were 585 alike–difference judgments and 576 completed semantic differential scales.

# Results

## Comparison of Percent Deidentification Between and Within Pairs

**Preliminary analysis of birth order effects.** Before proceeding with the study of variations in percent deidentification judgments between and within pairs, it was necessary to demonstrate that this percentage was independent of birth order within pairs. Table 2 shows the percent deidentification for each of the four pairs and for prior- and later-born

[1] Questions about parents were intended for a future study of the relationship between identification and deidentification.

members of each pair. Chi-square tests showed that none of the differences between prior- and later-born members within pairs was significant. For each pair, additional chi-square tests were undertaken to evaluate whether the lack of significant differences between prior- and later-borns held for each of the sample characteristics: sex of subject, male or female; sex of sibling, male or female and same or opposite sex; and spacing, narrow (below median year) or wide (above median year). None of these subtotals showed significant differences between prior- and later-borns. Finally, in addition to testing the effects of birth order within pairs, percent deidentification was compared for the combined judgments of first-, second-, and third-borns in the three-child family. Deidentification percentages were 60.6, 62.2, and 52,0, respectively, with no significant effect of birth order.

The lack of significant birth order effects provided the required justification for combining the judgments of both members of each pair in order to undertake the study of pairs.

**Within-pair effects.** For each of the four pairs, chi-square tests were undertaken comparing differences in percent deidentification for each of the within-pair sample characteristics: sex of subject, sex of sibling, and spacing (years). Only one pair, the first of the three-child family, showed a significant within-pair difference for only one sample characteristic, same- versus opposite-sex sibling. For this pair, percent deidentification was significantly higher for same-sex siblings, 84.6, than for opposite, 65.8, $\chi^2(1) = 7.34, p < .01$. None of the other pairs or sample characteristics showed significant within-pair differences.

In addition, because spacing measured in months is a continuous variable, it was possible to carry out three-way analyses of variance (Alike–Difference Judgment × Sex of Subject × Sex of Sibling) to further study the within-pair relationship between spacing and deidentification. Analyses of variance were undertaken for each pair and each member of each pair. In no analysis was the relationship between spacing and alike–difference judgment found to be significant, nor was there any significant relationship between age and deidentification for any sex-of-subject–sex-of-sibling combination.

Finally, since there is the persistent suggestion in the child psychology literature (e.g., Mussen, Conger, & Kagan, 1974, p. 443) that the 2- to 4-year spacing is especially difficult (Koch's, 1955, 1956a, 1956b, data usually form the basis for this impression), we proceeded to compare percent deidentification for 2- to 4-year spacing with that below 2 and above 4. None of the chi-squares for any of the four pairs was significant.

**Between-pair effects.** Table 2 shows the percent deidentification for each of the four pairs, 75.3 for the first pair of the three-child family; 52.7 for the second; 45.3 for the jump; and 62.2 for the first pair of the two-child family. The overall chi-square, comparing these four percentages was significant, $\chi^2(3) = 29.84, p < .001$, so that we proceeded with the comparison between first, second, and jump pairs. For the three-child family, Table 3 shows the results of chi-square tests between pair totals and between pair subtotals for each of the sample characteristics: sex of subject, sex of sibling, and spacing (years). Table 4 shows comparable chi-squares testing the differ-

**Table 2: Percent Deidentification for Pairs and for Prior- and Later-Born Members Within Pairs**

| Pairs and members | Percent deidentification | *n* |
|---|---|---|
| Three-child pairs | | |
| First pair | | |
| S1Sb2 (Prior) | 79.0 | 90 |
| S2Sb1 (Later) | 70.3 | 64 |
| Total | 75.3 | 154 |
| Second pair | | |
| S2Sb3 (Prior) | 52.4 | 63[a] |
| S3Sb2 (Later) | 53.1 | 49 |
| Total | 52.7 | 112 |
| Jump pair | | |
| S1Sb3 (Prior) | 42.2 | 90 |
| S3Sb1 (Later) | 51.0 | 49 |
| Total | 45.3 | 139 |
| Two-child pairs | | |
| S1Sb2 (Prior) | 60.0 | 100 |
| S2Sb1 (Later) | 65.0 | 80 |
| Total | 62.2 | 180 |

*Note.* S1, S2, S3 = first-, second-, and third-born subjects, respectively; Sb1, Sb2, Sb3 = first-, second-, and third-born siblings, respectively.
[a] Missing one second born's judgment of third-born sibling.

ences between the first pair of the two-child family and the second and jump pairs of the three-child family.

It can be seen in Tables 3 and 4 that both three-child and two-child first pairs showed significantly higher total percent deidentification than the jump pair. For the first pair of the three-child family (Table 3) this increment was significant for all of the sample characteristic subtotals. For the two-child family all of the subtotals showed increments relative to the jump pair, but some failed to reach significance for one of the two subtotals of a given characteristic, as shown in Table 4. Male subjects showed a significant increment, but not females; female siblings, but not males; opposite-sex siblings, but not same-sex; wide spacing, but not narrow. These discrepancies between subtotals for a given characteristic suggested a possible statistical interaction between type of pair and sample characteristic. Following Blalock (1972, p. 309), the significance of the interaction was tested by calculating $z$ for the difference between the difference between percent deidentification for the two subtotals of any sample characteristics showing this discrepancy. None of the interactions were found to be significant.

**Table 3: Deidentification Comparison Between Three-Child Pairs: Pair Totals and Sample Characteristics Subtotals**

| Total and subtotals | Pair | | |
|---|---|---|---|
| | First vs. Second | First vs. Jump | Second vs. Jump |
| Pair totals | 13.70* | 27.66* | 1.68 |
| Sample characteristics subtotals | | | |
| Sex of subject | | | |
| Male | 3.66 | 8.10* | .74 |
| Female | 10.42* | 20.30* | .97 |
| Sex of sibling | | | |
| Male | 2.50 | 12.92* | 3.49 |
| Female | 14.08* | 14.86* | .03 |
| Same | 11.20* | 23.04* | 2.22 |
| Opposite | 3.79 | 6.46* | .12 |
| Spacing (years)[a] | | | |
| Narrow | 7.93* | 10.26* | .26 |
| Wide | 7.24* | 6.26* | 1.85 |

*Note.* Data are derived from chi-square tests.
[a] Between-pair spacing dichotomy was obtained by maximizing the number of judgments in the cell with the smallest number, 3, 4, and 5 years for columns 1, 2, and 3, respectively.
* $p < .01$. (No differences were significant at .05 level.)

**Table 4: Deidentification Comparisons Between Two-Child First and Three-Child Second and Jump Pairs: Pair Totals and Sample Characteristics Subtotals**

| Total and subtotals | Pair | |
|---|---|---|
| | First vs. Second | First vs. Jump |
| Pair totals | 2.13 | 9.04** |
| Sample characteristics subtotals | | |
| Sex of subject | | |
| Male | 2.32 | 6.29** |
| Female | .37 | 3.18 |
| Sex of sibling | | |
| Male | .04 | 3.22 |
| Female | 5.70** | 5.85** |
| Same | .05 | 2.12 |
| Opposite | 5.17* | 8.65** |
| Spacing (years)[a] | | |
| Narrow | 1.46 | 2.84 |
| Wide | 1.47 | 5.08* |

*Note.* Data are derived from chi-square tests.
[a] Between-pair spacing dichotomy was obtained by maximizing the number of judgments in the cell with the smallest number, 4 and 5 years for columns 1 and 2, respectively.
* $p < .05$.
** $p < .01$.

That is, both first pairs showed significantly more deidentification than the jump pair, with no significant interactions with sample characteristics.

When the first pairs are compared with the second pair, the chi-square tests of Tables 3 and 4 show that deidentification is significantly higher for the total of the first pair of the three-child family, but not for the total of the two-child first pair. For the first pair of the three-child family, some of the subtotal increments fail to reach significance (i.e., male subject, male sibling, and opposite-sex sibling; see Table 3), but none of the tests of the significance of the interaction between type of pair and sample characteristic were significant. Interaction effects were also evaluated for subtotals comparing the first pair of the two-child family with the second pair where one of the two subtotals showed a significant increment for the first pair (i.e., female sibling and opposite-sex sibling, see Table 4). Again the interactions were not found to be significant.

The more consistent superiority of the first pair of the three-child family relative to the first of the two-child when each was compared to second and jump pairs promp-

ted a statistical comparison between the two first pairs. Here tests showed a significant increment for the three-child first pair total, $\chi^2(1) = 6.58$, $p < .01$, relative to the two-child first pair, as well as a significant interaction between same- versus opposite-sex sibling ($z = 2.34$, $p < .05$) such that the increment applied only to same-sex siblings. In reporting within-pair effects, we have already noted the significantly higher percent deidentification for the same-sex siblings of the first pair of the three-child family as compared to the opposite sex. This very high scoring same-sex pair with 84.6% deidentification seems to account for the more consistent superiority of the first pair of the three-child family relative to the first pair of the two-child when each is compared to second and jump pairs.

Another possible explanation for the more consistent picture for three-child pairs, relative to two-child pairs, was a methodological one. For each between-pair comparison, there is a set of four possible comparisons based on the two types of member judgments within each pair. For example, for the comparison between first and second pairs the following four types of member comparisons are involved: (a) S1Sb2 vs. S2Sb3, (b) S1Sb2 vs. S3Sb2, (c) S2Sb1 vs. S2Sb3, and (d) S2Sb1 vs. S3Sb2. It can be seen that for each between-pair comparison of three-child pair totals, one of these four possible combinations involves the same-member subject (S2Sb1 vs. S2Sb3 in the example above), while the other three are based on independent subjects. On the other hand, all four possible comparisons of the two-child pairs with the three-child, are based on independent-member subjects. To evaluate the influence of this same-subject comparison on the results for the three-child pairs, separate chi-squares were calculated for each of the four possible combinations of judgments per pair. (Chi-square tests for correlated proportions were applied for the same-subject comparisons, McNemar, 1962, p. 225.) Of the eight chi-squares comparing the first pair either with the second or with the jump pair, six were significant beyond the .05 level and two at the .06 (S2Sb1 vs. S3Sb2) and .07 (S2Sb1 vs. S2Sb3) levels. Of the latter two, one was an independent-subject comparison, while the other was a same-subject comparison, and both so closely approached significance that it would appear that any differences in the consistency of findings between first pairs of two- and three-child families cannot be attributed to this methodological consideration. Rather, the relative superiority of the first pair of the three-child family seems most parsimoniously explained by the very high levels of percent deidentification for the same-sex siblings of this pair.

The comparison between second and jump pairs (Table 3) shows no significant difference in percent deidentification for pair totals, nor for pair subtotals. (Nor were any of the four types of member comparisons significant.) Despite this lack of significance, the jump pair clearly shows the lowest levels of deidentification when both second and jump pairs are compared with first pairs. Both first pair totals are significantly higher than the jump pair totals, while only the three-child first pair is significantly higher than the second pair.

**Summary.** Significant within-pair effects were limited to the first pair of the three-child family and to the sample characteristic same- versus opposite-sex sibling, with the former showing higher deidentification. Significant between-pair effects show that first pairs produce the highest levels of deidentification, jump pairs the lowest. There is also some evidence that second pairs deidentify at intermediate levels. Finally, none of the comparisons between first, second, and jump pairs showed significant interactions with sample characteristics.

## Comparison of Alike–Difference Judgments and Semantic Differentials

**P score.**[2] Nonparametric statistics were applied in studying the P score, because the distribution was both skewed and leptokutic. The median number of polarized items was 2.91 with a range of 0 to 13.

[2] Rank correlation between the D score of Osgood and Suci (1952) and the P score was found to be .82 (F. F. Schachter, Note 1), suggesting that polarized differences account for 67% of the variance in D scores.

The tests show that the association between the P score and the global alike–difference judgment was significant. Median tests yield chi-squares of 23.79, if we dichotomize the P score at two items polarized, and 100.49, if we dichotomize at three items polarized ($p < .001$, $df = 1$ for both). The mean number of polarized items was 1.89 for subjects judging their siblings alike and 3.50 for those judging different, with the Mann-Whitney $U$ significant at the .001 level.

In addition, the P score yielded between-pair results that closely matched those for the alike–difference judgment. It will be recalled (see Table 2) that the latter showed 75.3% deidentification for the first pair of the three-child family, 62.2% for the first pair of the two-child family, 52.7% for the second pair and 45.3% for the jump pair. The P score, dichotomizing at 3.0 (as near the median of 2.91 as possible following McNemar, 1962, p. 376) showed the same relationships between pairs, 43.0% with more than 3.0 polarized items for the first pair of the three-child family, 37.4% for the first pair of the two-child family, 27.9% for the second pair, and 25.9% for the jump pair. As with the global judgments, the overall comparison among the four pairs was significant, median test $\chi^2(3) = 12.09, p < .01$, as well as the differences between both of the first pairs and the jump pair, $\chi^2(1) = 9.19, p < .01$ for the three-child family, $\chi^2(1) = 4.64, p < .05$ for the two-child family; and also the increment for the first pair of the three-child family relative to the second pair was significant, $\chi^2(1) = 6.30, p < .05$.

**Item analysis.** In view of the significant association between global judgments and total polarized items, we proceeded with the item analysis. All 13 items showed more polarization when subjects judged their siblings to be different rather than alike. For 11 of the 13 items, except only tense–relaxed and rugged–delicate, the increment for different-judging subjects was significant beyond the .001 level. Looked at from the perspective of the frequency distribution of polarization per item, the items tense–relaxed and rugged–delicate are the ones that showed relatively more polarization in the distribution for alike-judging subjects compared to that for different-judging subjects. The chi-square comparison of the two distributions was 26.21, $p < .05$, $df = 12$, with only the items tense–relaxed and rugged–delicate showing significant differences. The other items were equally represented in both distributions.

Frequency distributions of polarizations per item were also examined to see if different judgments of opposite-sex siblings were more likely to be associated with polarization on sex-stereotyped dimensions than were those of same-sex siblings. Chi-squares showed that the frequency distributions of polarizations per item, when subjects were judging same-sex siblings as compared to opposite-sex, differed insignificantly, regardless of whether the subject's global judgment was alike or different. Nor were the frequency distributions significantly different when judgments of male and female subjects were compared.

## Discussion

### Possible Determinants of Deidentification

On the basis of our results, we shall consider possible determinants of sibling deidentification in the light of the existing psychoanalytic, social learning, and cognitive–developmental theories of identification, as well as Bossard and Boll's (1956) theory of role differentiation among siblings.

Applying the psychoanalytic framework, the finding that deidentification was highest in first pairs, lowest in jump pairs, and intermediate in second pairs suggests the hypothesis that deidentification may be a defense mechanism for mitigating sibling rivalry, a Cain Complex, just as identification is regarded as a mechanism for resolving child–parent rivalry, the Oedipus Complex. If deidentification were such a mechanism, the factors that are likely to increase sibling rivalry would be expected to increase the level of deidentification. It seems likely that the rivalry of the first pair is more intense than that of the jump pair since the rivalry between jump pair members would probably be mitigated by the presence of the intervening sibling. First-pair rivalry would

be expected to be highest, undiluted by the inevitable delay in arrival of the third born. Second-pair rivalry would likely be at intermediate levels, unmitigated by an intervening sibling as in the jump pair, yet somewhat diluted by the perpetual presence of a third sibling. It can be seen that the significant between-pair differences in deidentification correspond with expected variations in sibling rivalry, supporting the rivalry–defense hypothesis.

The within-pair finding that same-sex siblings are significantly more likely to deidentify than are opposite-sex siblings, though limited to the first pair of the three-child family, also supports the psychoanalytic hypothesis. The shared desires of same-sex siblings provide a basis for greater rivalry than would be the case for opposite-sex siblings, so that higher levels of defensive deidentification would be expected with more intense rivalry.

The fact that subjects who judge their siblings alike, compared to those who judge their siblings different, are relatively more likely to think of themselves and their siblings in terms of the dimension tense–relaxed also supports the rivalry–defense hypothesis. Defenses should operate to reduce tension so that alike-judging subjects, who may not be resorting to the defense of deidentification, may experience the tense–relaxed dimension with greater salience, relative to different-judging subjects.

Although the rivalry–defense hypothesis seems useful in explaining much of the present data, one might ask why the defense of deidentification, why not sibling identification as with child–parent rivalry? Here psychoanalytic theory provides a reasonable answer in that identification is predicated on the postponement of gratification to adolescence. That is, the Oedipal child says, in effect, "When I grow up I'll marry someone like my opposite-sex parent." With siblings growing up together there is no possibility for mitigating rivalry by postponement. Indeed, identification would only serve to intensify sibling rivalry, both siblings seeking the same gratifications at approximately the same time.

Another question with regard to the rivalry–defense hypothesis concerns the finding of no significant effect of spacing on sibling deidentification. On the basis of the sibling rivalry literature one might expect a 2–4-year spacing to generate greater rivalry than spacing less than 2 or greater than 4 years. Yet we found no significant spacing effects on deidentification within-pairs and no significant interaction between type of pair and spacing in the between-pair study. This finding prompted a reexamination of Koch's data (1955, 1956a, 1956b), since her sample seems to form the basis for existing generalizations on spacing and rivalry. Close scrutiny fails to support the oft-quoted conclusion that 2–4-year spacing generates the most rivalry. On the one hand, Koch's subjects were merely 5 or 6 years old, so that siblings 4–6 years younger would be newborns, confounding effects of spacing with effects of acquiring a new sibling. On the other hand, Koch's findings on spacing are endlessly qualified by considerations of both sex and birth order, so that she concludes one of her later reports as follows: "Since the spacing variable was related in a very complex fashion to our other variables, we shall attempt no description here of the web of relations" (Koch, 1956b, p. 424). Altogether, Koch's data are ambiguous enough to leave quite open the question of spacing effects on sibling rivalry. If so, then the findings of no significant spacing effects on deidentification need not represent a problem for our rivalry–defense hypothesis. Indeed, the present findings, in conjunction with the fact that jump pair spacing is so much wider than first pair (see Table 1), suggest that some of the generalizations concerning spacing and sibling rivalry may have resulted from a confounding with pair effects.

With regard to social learning theory of identification (Bandura, 1962; Bandura & Huston, 1961; Bandura, Ross, & Ross, 1963), the concept that identification increases with greater similarity, power, and nurturance of the model does not seem to be supported by the present data. For three of the four pairs, similarity of sex (same- versus opposite-sex sibling) has no significant effect on the alike–difference judgment. For the one pair, where similarity of sex does have an effect, in the case of the first pair of

the three-child family, the effect is contrary to the prediction from social learning theory, since same-sex siblings deidentify more than do opposite-sex siblings for this pair. With regard to power, one would expect younger siblings to identify more with older siblings especially when spacing is wide; since a much older sibling should be viewed as more powerful. Yet the data show no significant effects of birth order or spacing on alike–difference judgments and no significant interaction. Nor does nurturance of the model seem to effect sibling deidentification. One might expect older sisters to be the most nurturant group; yet we found no significant interaction between birth order and sex of subject or sibling with regard to deidentification. It may be that other measures of being like or different from siblings will provide support for social learning theory, but the present data on self-perceived likeness provide little support.

Kohlberg's (1966) cognitive–developmental theory of identification is useful in suggesting, by analogy with his discussion of the development of cognitive categories for male versus female, that the categories alike versus different are probably not applied to siblings before the age when stable concepts of alike versus different are organized cognitively. However, the theory offers little help in explaining why these judgments are applied to siblings in the first place, and why judgments of different predominate for first pairs. It goes without saying that the developmental course of deidentification requires investigation, not solely to assess age of onset, but also to see whether the phenomenon occurs at other than the college-age level.

With regard to Bossard and Boll's (1956, chap. 10) theory that siblings in a large family seek personality role differentiation to garner some distinctive recognition in a crowded household, the finding of more deidentification in the two-child first pair than in the second and third pairs of the three-child family fails to support their thesis. Even more undermining of Bossard and Boll's thesis is the fact that members of first pairs in the three-child family strive for distinction from each other significantly more often than do members of second and jump pairs. Their theory would imply an even distribution of differentiation between pairs of siblings, each member in search of some means to distinguish himself from the other.

On the other hand, a combination of the hypothesis of Bossard and Boll (1956), that larger families generate more sibling differentiation, with the hypothesis of rivalry–defense provides a plausible explanation for the fact that same-sex first pairs in the three-child family deidentify significantly more often than in the two-child family. That is, larger families may serve to intensify existing high levels of rivalry generated by same-sex first-pair competition for shared goals, thereby magnifying the degree of deidentification.

### Meaning of Global Judgments

The results of the second study indicate that when subjects judge their siblings to be different from themselves they mean opposite on a broad range of personality dimensions. All but two of the 13 bipolar scales of the semantic differential showed significantly more polarization for the subjects judging their siblings different than for those judging alike, and the frequency distribution of item polarization was the same whether subjects were judging same- or opposite-sex siblings, whether subjects were male or female. Of the two items that did not discriminate between subjects judging alike and different, rugged–delicate and tense–relaxed, the latter has already been discussed in connection with the psychoanalytic hypothesis (see above). As for the former, it may have been neglected as a basis for judging different because of our instructions to judge siblings on the basis of personality characteristics. Rugged–delicate may have been viewed as a physical characteristic.

Before concluding, a comment concerning the lack of birth order effect within pairs on deidentification seems warranted, in view of the fact that most of our impressions on the nature of firstborns compared to later borns are based on comparisons with second borns (Koch, 1955, 1956a; 1956b; Sutton-Smith & Rosenberg, 1970; S. Schachter's

studies, 1959, chaps. 5–6, 1964, are a notable exception with a heterogeneous later-born group). If, as the present findings suggest, differences between firstborns and second borns are magnified by the process of polarization, generalizations concerning the nature of firstborns will need to be qualified by reference to the birth order of the comparison group.

In conclusion, the results support the psychoanalytic formulation that deidentification is a mechanism for resolving sibling rivalry, a Cain Complex. This is, of course, a first study and further evidence will be needed. Studies are indicated under conditions where sibling rivalry is likely to abate as, for example, with the mellowing of the mature years or with communal rather than nuclear-family upbringing. In addition, the phenomenon of deidentification itself (its developmental course, how parents judge their children, etc.) seems to merit further investigation, regardless of its determinants. Finally, deidentification may have clinical and educational implications. How many underachievers are deidentifying with achieving siblings, how many delinquents with conventional siblings? Such polarization may begin at an early age and form the basis of a self-fulfilling prophesy. Perhaps blatant rivalry is preferable to an early constriction in the range of possibilities of the self.

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