Developmental Psychology
1988, Vol. 24, No. 3, 386–392

0012-1649/88/$00.75

Mother–Infant Face-to-Face Interaction: Influence is Bidirectional and Unrelated to Periodic Cycles in Either Partner's Behavior

Jeffrey F. Cohn
University of Pittsburgh

Edward Z. Tronick
University of Massachusetts

During mother–infant face-to-face interactions, bidirectional influence could be achieved through either the entraining of periodic cycles in the behavior of each partner or through the stochastic organization of behaviors. To determine whether and how bidirectional influence occurs, we used both time- and frequency-domain techniques to study the interactions of 54 mother–infant pairs, 18 each at 3, 6, and 9 months of age. Behavioral descriptors for each mother and infant were scaled to reflect levels of affective involvement during each second of the interaction. Periodic cycles were found in infants' expressive behavior only at 3 months and not in mothers' behavior. Nonperiodic cycles, which were found in some mothers' and infants' behavior at each age, were more common. At no age was the occurrence of cycles in mothers' or infants' behavior related to the achievement of bidirectional influence. Similar proportions of mothers and infants were responsive to moment-to-moment changes in the other's behavior, except at 6 months when the proportion of mothers was higher. Bidirectional influence was brought about by the stochastic organization of behaviors rather than through the mutual entraining of periodic cycles.

Early mother–infant face-to-face interactions have a conversation-like pattern in which each partner appears to be responsive to the other. The assumption that this pattern is actually achieved by bidirectional influence has been seriously questioned in a series of papers (Gottman & Ringland, 1981; Thomas & Malone, 1979; Thomas & Martin, 1976). Few studies have rigorously tested the null hypothesis that during face-to-face interactions moment-to-moment changes in the infant's behavior are independent of changes in the mother's behavior. Three studies that did test the null hypothesis (Gottman & Ringland, 1981; Hayes, 1984; Thomas & Malone, 1979) failed to reject it.

Two types of organization of the infants' behavior, periodic or stochastic, would permit the mother to create the semblance of bidirectional influence. *Periodic* events cycle on and off at regular, precise intervals, permitting highly accurate prediction of the timing of future events. A periodic cycle is deterministic in that the frequency, phase, and amplitude do not vary over time (Gottman, 1981). Alternatively, *stochastic* events are autocorrelated over short intervals; that is, sequences occur nonrandomly (e.g., smiles following the onset of visual regard; Kaye & Fogel, 1980). Depending on the type of autocorrelation, sequences may also be cyclic, but not periodic. Cohn and Tronick (1983), for instance, reported that, during normal interactions, infants displayed cycles of neutral and positive expressions. Because these cycles were stochastic (i.e., nonperiodic and, hence, variable in frequency, phase, and amplitude), they would not accurately predict infants' expressions over the long term.

These two types of behavioral organization have different implications for how bidirectional influence could occur. One hypothesis is that periodic cycles occur in both the infant's and the mother's behavior and that these cycles become synchronized through a process of mutual entrainment (Lester, Hoffman, & Brazelton, 1985; Schaffer, 1977). A second hypothesis is that expressive behaviors are autocorrelated over short intervals but also cross-correlated with (i.e., contingent on) the preceding behavior of the partner (Cohn & Tronick, 1987; Kaye & Fogel, 1980). These hypotheses are not mutually exclusive; a third hypothesis, therefore, is that bidirectional influence occurs in both of these ways.

These hypotheses all posit active processing of social signals by the young infant and are consistent with the view that infants respond in specific and appropriate ways to their mother's communicative displays (Campos, Barrett, Lamb, Goldsmith, & Stenberg, 1983; Cohn & Tronick, 1983; Tronick, 1981). They differ in regard to how this responsiveness comes about.

Mutual entrainment hypotheses make greater demands on the infant's cognitive abilities because they assume that the infant can abstract relatively long periodicities from the mother's behavior. Lester, Hoffman, and Brazelton (1985) reported periodicities of 10 to 45 s. The accomplishment of this task would be all the more impressive in light of the fact that the mother's periodicity varies as she attempts to adjust to her infant's. Because face-to-face interactions seldom last more than several minutes, infants might have to accumulate experience over the course of many interactions before mutual entrainment could occur and bidirectional influence could be detected.

Preparation of this article was supported in part by National Institute of Mental Health Grant 1R03MH39706-01A1 to the University of Pittsburgh.

Elizabeth Krafchuk, Margaret Ricks, and Steve Winn assisted with the collection of these data, and Reinaldo Matias and Shelley Ross assisted with data analysis. David Stoffer provided statistical consultation.

Correspondence concerning this article should be addressed to Jeffrey F. Cohn, Psychology Department, Clinical Psychology Program, 4015 O'Hara Street, University of Pittsburgh, Pittsburgh, Pennsylvania 15260.

Data relevant to these hypotheses have been inconsistent. Brazelton, Koslowski, and Main (1974) described infants' attention to the mother as cycling on and off at four regular intervals per minute and hypothesized that these cycles were sinusoid, which implies periodicity. Nevertheless, they also claimed that the duration of cycles varied depending on the quality of the mother's behavior. Lester, Hoffman, and Brazelton (1985), on the other hand, appear to have assumed that any nonrandom temporal organization is periodic. Visual analysis of their graphs of mothers' and infants' spectral density functions suggests otherwise. Such graphs are characteristic of stochastic organization (see Appendix A for an example).

The present report uses both time- and frequency-domain techniques (Gottman, 1981; McCleary & Hay, 1980) with mother–infant pairs at 3, 6, and 9 months to define the organization of infants' behavior. Three months was chosen because previous theory and some data about periodicity and bidirectional effects have focused on infants at this age (Brazelton et al., 1974; Kaye & Fogel, 1980; Gottman & Ringland, 1981; Stern, 1974). Nine months was chosen because recent data (Jasnow & Feldstein, 1986; Martin, 1981) suggest that by 9 or 10 months infants' behavior during face-to-face interactions is stochastic and influence is bidirectional. The organization of behavior may change with development, and the range of ages studied allowed us to investigate that possibility.

Method

Subjects and Procedures. The subjects were 18 infants, 9 boys and 9 girls, each at 3, 6, and 9 months of age. All were from middle-class families and had experienced no significant perinatal medical complications. Face-to-face interactions were videotaped in our laboratory using a split-screen procedure. Only the first of three 2-min interactions are included in this report (additional details are reported in Cohn & Tronick, 1987).

Coding. Videotapes were coded using the Monadic Phases Manual (Tronick & Cohn, 1987), which is a revised version of a system described by Tronick, Als, and Brazelton (1980; Als, Tronick, & Brazelton, 1979). Following Tronick (Tronick, Als, & Brazelton, 1977) and Lester (Lester et al., 1985), the monadic phases were scaled along an attentional/affective dimension. For our coding of monadic phases, this resulted in a 9-point scale; Lester et al. (1985) used a more differentiated coding of monadic phases that resulted in a 13-point scale.

Mother and infant monadic phases were scored separately by teams of two coders using a stop-frame procedure: Whenever a change in phase was observed, the videotape would be reversed and replayed at full and at slow speed to determine whether a change and what type of change in phase had occurred and the time of its occurrence. Times, read from the digital time display, were rounded to the nearest 0.25 s. For comparison, videotapes of 12 mothers and 5 infants were recoded by a second team of coders. Agreement, defined as the second team of coders' observing the same phase within 0.5 s of the first, ranged from 81% to 97% for mothers' monadic phases and from 90% to 100% for infants' monadic phases (Ks = 60% and 72%, respectively).

Data reduction. To achieve comparability with previous research that used a 1-s modified frequency scoring interval (Gottman & Ringland, 1981; Lester et al., 1985; Tronick et al., 1977), scaled scores were averaged within 1-s blocks. This produced a time series for each mother and infant of approximately 120 observations, which is sufficiently long for time-series analysis. All analyses were of individual subjects.

Data analysis. To determine whether a stochastic or a periodic process characterized the time series for each mother and infant, and as a preliminary step in the analysis of bidirectional influence, each univariate time series was modeled using time-domain Autoregressive Integrated Moving Average (ARIMA) analysis (Appendix B). Unlike the frequency-domain spectral analysis used by Lester et al. (1985), ARIMA analysis provides (orthogonal) parameter estimates of stochastic and periodic components in a time series. The two approaches are theoretically equivalent, but in practice the time-domain approach may miss periodic processes that add little variance to a time series.

To guard against such errors, we computed spectral density functions (Appendix C) for each series and compared them with those expected on the basis of the fitted time-domain model. When a spectral analysis indicated a periodicity not present in the model, we refit the series with a periodic parameter at a lag consistent with the spectral analysis.

As an initial step in the analysis of bidirectional influence, we first computed cross-correlation functions (CCFs) from the raw (or first-differenced, as appropriate) data for each mother–infant pair. These results were used to screen out dyads for whom no further analyses were warranted. A finding of no statistically significant cross-correlations (i.e., all cross-correlations within two standard deviations of zero) is sufficient evidence to rule out bidirectional influence. Where significant cross-correlations were found, we used Gottman's (Williams & Gottman, 1982) program Bivar (Appendix D) to test the null hypothesis that mother's and infant's behaviors are independent.

Results

The time series for all but one of the mothers and one of the babies were stationary and did not require differencing. In the two cases that required differencing, first-order differencing was sufficient to bring about stationarity.

Is the behavior of mothers periodic? The time-domain analyses showed no evidence of periodicity in any mother's time series, with the exception of one mother each at 3 and at 6 months. The absence of periodicity was confirmed by spectral analysis.

Is the behavior of infants periodic? At 3 months, the time series for 5 of 18 (28%) babies had a significant periodic component. Of these, 3 were identified during the initial time-domain modeling; 2 others suggested periodicity after spectral analysis. The mean period was 10 s. The variance due to periodicity, however, was small relative to that accounted for by autoregression (see Is the Infant's Behavior Stochastic below). For these 5 babies, periodicity accounted for less than 3% of the variance in the univariate time series.

Only 1 of 18 infants at 6 months and none at 9 months showed any evidence of periodicity. Both of these proportions were below the 95% confidence interval for 3 months.

Is the mother's behavior stochastic? The mothers' series all had significant autoregressive parameters and, as noted before, no significant periodic parameters. Autoregression accounted for an average of 37% of the variance in the mothers' univariate time series; this proportion did not vary with age of the infant.

Four mothers at 3 months, 8 mothers at 6 months, and 6 mothers at 9 months had series that were fit by nonperiodic, cyclic AR(2) models. Estimated mean cycle durations were 16 s, 23 s, and 27 s at 3, 6, and 9 months, respectively. Neither the number of mothers with nonperiodic cyclic series nor the duration of the cycles varied with infant age.

Is the infant's behavior stochastic? The time series for all but one of the infants had significant autoregressive parameters. Autoregression accounted for an average of 36% of the variance

Table 1
Proportion of Mothers and Infants at Each Age Whose Univariate Time Series Included a Periodic or a Stochastic, Nonperiodic Cycle

	Type of cycle	
Dyad member	Periodic	Stochastic, nonperiodic
Mothers		
3 months	.06	.22
6 months	.06	.44
9 months	.00	.33
Infants		
3 months	.28	.17
6 months	.06	.44
9 months	.00	.22

Note. N = 18 mothers and 18 infants at each age. For all but 1 baby and no mother, the periodic cycle is in addition to a large stochastic component. The maximum number of cycles is, therefore, two: one stochastic and one periodic.

in infants' univariate time series; this percentage did not vary with age of infant nor did it differ from that for the mothers.

Nonperiodic, cyclic AR(2) models fit the series of 4 infants at 3 months, 8 infants at 6 months, and 4 infants at 9 months (an example from the 6-month data appears in Figure A1 in Appendix 1). Estimated mean cycle durations were 23 s, 18 s, and 17 s at 3, 6, and 9 months, respectively. Neither the number of babies with nonperiodic cycles nor the cycle durations varied with infant age. Table 1 summarizes these findings.

Are changes in the infant's behavior related to changes in the mother's behavior? The proportion of infants who responded to changes in their mother's behavior was similar to the proportion of mothers who responded to changes in their infant's behavior, except at 6 months (see Table 2). The size of bidirectional effects was also similar for both partners. When the bivariate models were significant, they increased the proportions of explained variance by 16% for mothers and 17% for infants.

Babies were more likely to respond to changes in their mother's behavior if the mother were responsive to changes in their behavior. Only one baby at each age showed a bidirectional effect when the mother did not.

To see whether the occurrence of cycles—either periodic or nonperiodic—in the infant's behavior was related to the achievement of bidirectional influence, we first compared the proportion of dyads in which both the mother's and the infant's behavior was cyclic with the proportion predicted by chance (i.e., we tested to see whether the occurrence of cycles in the mother's behavior was independent of the occurrence of cycles in the baby's behavior). Second, we compared the proportion of infants whose behavior was both cyclic and bidirectional with the proportion predicted by chance (i.e., we tested to see whether the occurrence of cycles in the baby's behavior and the achievement of bidirectional influence were independent). Both proportions were nonsignificant ($z = -.01$, and $z = -.43$, respectively).

Discussion

Mothers' and babies' time series, with few exceptions, were stationary. This finding is relevant to the generalizability of our other findings. Had we looked at longer interactions, of the length studied by Lester et al. (1985) or Gottman and Ringland (1981), would we have found otherwise? The available evidence suggests not. Gottman and Ringland analyzed three 3-min interactions and found nonstationarity, but only at the longest cycles. They reported that the effect was not large. Lester et al. analyzed eighty 3-min interactions and found no evidence of nonstationarity, although they had first removed any linear trend. Cohn and Elmore (1987) analyzed twenty 3-min interactions and found few exceptions to stationarity. Thus, only one study has found nonstationarity and, then, only at the slowest frequencies. Moreover, the convergence between our findings and those of Lester et al., despite differences in length of interaction and method of analysis, argues for the robustness of our findings with respect to stationarity.

We found strong support for the belief that face-to-face interactions are a product of bidirectional influence. At 3 and 9 months, similar proportions of infants and mothers had significant bivariate models. At these ages, we also found that infants and mothers were equally influential in influencing the direction of the interaction. At 6 months, on the other hand, mothers were more likely to follow their infant's lead. The relative difference between mothers and babies at 6 months may be related to infants' increased interest in objects at this age (Cohn & Tronick, 1987).

Some previous studies that found bidirectional influence in mother–infant interactions have been seriously criticized on statistical grounds. Rigorous reanalyses of data from Jaffe, Stern, and Peery (1973) and from Tronick, Als, and Brazelton (1977) failed to replicate the original findings (Gottman & Ringland, 1981; Thomas & Malone, 1979). A study by Hayes and Elliot (reported in Hayes, 1984) found that mothers' and infants' vocalizations were independent.

Two recent studies, in addition to ours, have used appropriate analyses and found evidence of bidirectional influence (Beebe, Jaffe, Feldstein, Mays, & Alson, 1985; Lester et al., 1985). An important difference between these studies and previous ones is that they regarded expressive behavior as multimodal. Jaffe et al. (1973) and Hayes and Elliot (reported in Hayes, 1984) examined only individual response modalities (e.g., gaze or vocalization) or small numbers of interactions. Under these conditions, evidence of bidirectional influence may be less likely.

Several factors demonstrate that bidirectional influence was achieved through the stochastic organization of behaviors rather than through the mutual entrainment of periodicities. First, if periodicity were, as Lester et al. (1985) and others argue,

Table 2
Proportion of Mothers and Infants at Each Age Who Were Responsive to Changes in the Other's Behavior

Infant's age	Mothers	Infants	z
3 months	.55	.39	*ns*
6 months	.67	.33	2.03
9 months	.50	.39	*ns*

Note. N = 18 mothers and 18 infants at each age; z is the test statistic for the significance of the difference between two proportions.

a biologically based "fundamental property of early face-to-face mother–infant interactions" (pp 22–25) we would have found a higher proportion of truly periodic cycles. We observed 54 mother–infant interactions and found periodic cycles in the time series of only 6 babies and 2 mothers, and even then periodicity accounted for little variance. It is possible, of course, that had we observed longer episodes of interaction we would have found a higher proportion of periodic cycles. However, the cycle durations that we observed were within the range reported by others (Brazelton et al., 1974; Lester et al., 1985).

Second, even if we were not to reject the mutual entrainment hypothesis on the grounds that too few infants or mothers showed truly periodic cycles, the total proportion of cycles, both periodic and nonperiodic, was still less than one would expect on the basis of Lester et al.'s characterization. No more than half of the babies showed any evidence of cycles, periodic or nonperiodic, at any age. The same was true of the mothers. Proponents of periodicity in infants' expressive behavior claim that it is always present. Our results clearly do not support that claim.

Third, even if the type and proportion of cycles that we found were consistent with the mutual entrainment hypothesis, the occurrence of cycles in the infant's behavior was independent of the occurrence of cycles in the mother's behavior. Furthermore, mothers or infants with cyclic behavior were no more likely to show responsiveness to their partner. Cyclic behavior and bidirectional influence were unrelated.

Our conclusions differ from those of Lester et al. but are consistent with their spectral data. Visual analysis of their spectral plots, as noted earlier, suggests stochasticity rather than periodicity, and our data confirm that impression.

A minor difference between our findings and theirs is that we found no more than two cycles (periodic plus stochastic) in any one time series. They identified three. This difference may be due to their practice of arbitrarily defining *slow, medium,* and *fast* cycles and averaging spectral densities within these ranges. This procedure may have overestimated the number of cycles, because one band of elevated frequencies, as occurs in AR(1) and AR(2) processes, could be counted as more than one "cycle." Moreover, they used a 13-point scale, which may have resulted in the identification of additional cycles. In this connection, Lester et al. reported that the use of fewer scale-points eliminated slower frequency cycles but did not otherwise influence the spectra.

Our findings seriously call into question the belief that infants' behavior during face-to-face interactions is periodic. In light of the results, spectral data previously cited in support of the periodicity hypothesis are better interpreted as indicative of stochastic organization. Of course, it is possible that were infants' behavior described with a system other than the monadic phases (as we use or as used by Lester et al. and Gottman & Ringland) or were some discrete behavior, such as vocalization or gaze, studied, acceptable evidence of periodicity might be found. Further research is needed to explore this possibility.

The results provide an important basis for the use of sequential analysis to study mother–infant interactions. Sequential analysis is valid only if behavior is generated by a stationary, stochastic process. Were infants' behavior during interactions either nonstationary or periodic, its characterization by sequential analysis—and, consequently, findings from a large number of studies (e.g., Cohn & Tronick, 1983, 1987; Kaye & Fogel, 1980; Malatesta & Haviland, 1982; Stern, 1974)—would be potentially invalid.

In summary, using rigorous data analytic techniques, we have confirmed that the conversation-like pattern of mother–infant face-to-face interactions at 3, 6, and 9 months is produced by bidirectional influence. Bidirectional influence is achieved through the stochastic organization of behaviors and not through mutual entrainment of periodic cycles. Periodic cycles accounted for some variance in infants' behavior at 3 months. Nonperiodic cycles were found in some mothers' and infants' behavior at each age. However, at no age was the occurrence of cycles in mothers' or infants' behavior related to the achievement of bidirectional influence.

References

Als, H., Tronick, E., & Brazelton, T. B. (1979). Analysis of face-to-face interaction in infant–adult dyads. In M. E. Lamb, S. J. Suomi, & G. R. Stephenson (Eds.), *Social interaction analysis: Methodological issues* (pp. 33–76). Madison: University of Wisconsin.

Beebe, B., Jaffe, J., Feldstein, S., Mays, K., & Alson, D. (1985). Interpersonal timing: The application of an adult dialogue model to mother–infant vocal and kinesic interactions. In T. M. Field & N. A. Fox (Eds.), *Social perception in infants* (pp. 217–248). Norwood, NJ: Ablex.

Brazelton, T. B., Koslowski, B., & Main, M. (1974). The origins of reciprocity: The early infant–mother interaction. In M. Lewis & L. A. Rosenblum (Eds.), *The effect of the infant on its caregiver* (pp. 49–76). New York: Wiley.

Campos, J. J., Barrett, K. C., Lamb, M. E., Goldsmith, H., & Stenberg, C. (1983). Socioemotional development. In P. H. Mussen (Series Ed.) and J. J. Campos & M. H. Haith (Vol. Eds.), *Handbook of child psychology: 2. Infancy and developmental psychology* (pp. 783–916). New York: Wiley.

Chatfield, C. (1980). *The analysis of time series: An introduction.* London: Chapman and Hall.

Cohn, J. F., & Elmore, M. (1987). [Three-month-old infants' response to 5 s of mother's still face contingent on their becoming positive]. Unpublished raw data.

Cohn, J. F., & Tronick, E. Z. (1983). Three-month-old infants' reaction to simulated maternal depression. *Child Development, 54,* 185–193.

Cohn, J. F., & Tronick, E. Z. (1987). Mother–infant face-to-face interaction: The sequencing of dyadic states at 3, 6, and 9 months. *Developmental Psychology, 23,* 68–77.

Gottman, J. M. (1981). *Time series analysis: A comprehensive introduction for social scientists.* New York: Cambridge University Press.

Gottman, J. M., & Ringland, J. T. (1981). The analysis of dominance and bidirectionality in social development. *Child Development, 52,* 393–412.

Hayes, A. (1984). Interaction, engagement, and the origins and growth of communication: Some constructive concerns. In L. Feagans, C. Garvey, & R. Golinkoff (Eds.), *The origins and growth of communication* (pp. 136–161). Norwood, NJ: Ablex.

Jaffe, J., Stern, D., & Peery, C. (1973). "Conversational" coupling of gaze behavior in prelinguistic human development. *Journal of Psycholinguistic Research, 2,* 321–329.

Jasnow, M., & Feldstein, S. (1986). Adult-like temporal characteristics of mother–infant vocal interactions. *Child Development, 57,* 754–761.

Kaye, K., & Fogel, A. (1980). The temporal structure of face-to-face communication between mothers and infants. *Developmental Psychology, 16,* 454–464.

Lester, B., Hoffman, J., & Brazelton, T. B. (1985). The rhythmic structure of mother–infant interaction in term and preterm infants. *Child Development, 56,* 15–27.

Malatesta, C. Z., & Haviland, J. M. (1982). Learning display rules: The socialization of emotional expression in infancy. *Child Development, 53,* 991–1003.

Martin, J. (1981). A longitudinal study of the consequences of early mother–infant interaction: A microanalytic approach. *Monographs of the Society for Research in Child Development, 46* (Serial No. 190).

McCleary, R., & Hay, R. A. (1980). *Applied time series analysis for the social sciences.* Beverly Hills, CA: Sage.

Ryan, T. A., Joiner, B. L., & Ryan, B. F. (1982). *Minitab reference manual* (Minitab Project). University Park: Pennsylvania State University, Department of Statistics.

Schaffer, H. R. (Ed.). (1977). *Studies in mother–infant interaction.* New York: Academic.

Stern, D. (1974). Mother and infant at play: The dyadic interaction involving facial, vocal, and gaze behaviors. In M. Lewis & L. A. Rosenblum (Eds.), *The effect of the infant on its caregiver.* New York: Wiley.

Thomas, E. A. C., & Malone, T. W. (1979). On the dynamics of two-person interactions. *Psychological Review, 86,* 331–360.

Thomas, E. A. C., & Martin, J. A. (1976). Analyses of parent–infant interaction. *Psychological Review, 83,* 141–156.

Tronick, E. (1981). Infant's communicative intent. In B. Stark (Ed.), *Language behavior in infancy and early childhood* (pp. 5–16). Holland: Elsevier.

Tronick, E., Als, H., & Brazelton, T. B. (1977). Mutuality in mother–infant interaction. *Journal of Communication, 27,* 74–79.

Tronick, E., Als, H., & Brazelton, T. B. (1980). Monadic phases: A structural descriptive analysis of infant-mother face-to-face interaction. *Merrill-Palmer Quarterly of Behavior and Development, 26,* 3–24.

Tronick, E. Z., & Cohn, J. F. (1987). *Revised Monadic Phases Manual.* Unpublished manuscript.

Williams, E. A., & Gottman, J. (1982). *A user's guide to the Gottman-Williams time series analysis computer programs for social scientists.* New York: Oxford University Press.

Appendix A

Different temporal organizations have characteristic spectral profiles. A stochastic second-order autoregressive organization—AR(2)—has one nonperiodic cycle when the two autoregressive parameter estimates meet the following condition: $a_1^2 + 4a_2 < 0$. The spectral density function will peak over a broad range of frequencies, which reflects the variability of the cycle. Figure A1 is an example from our data of an autoregressive organization of this type. In this example, $a_1 = .75$ and $a_2 = -.24$; therefore, $a_1^2 + 4a_2 = -0.40$.

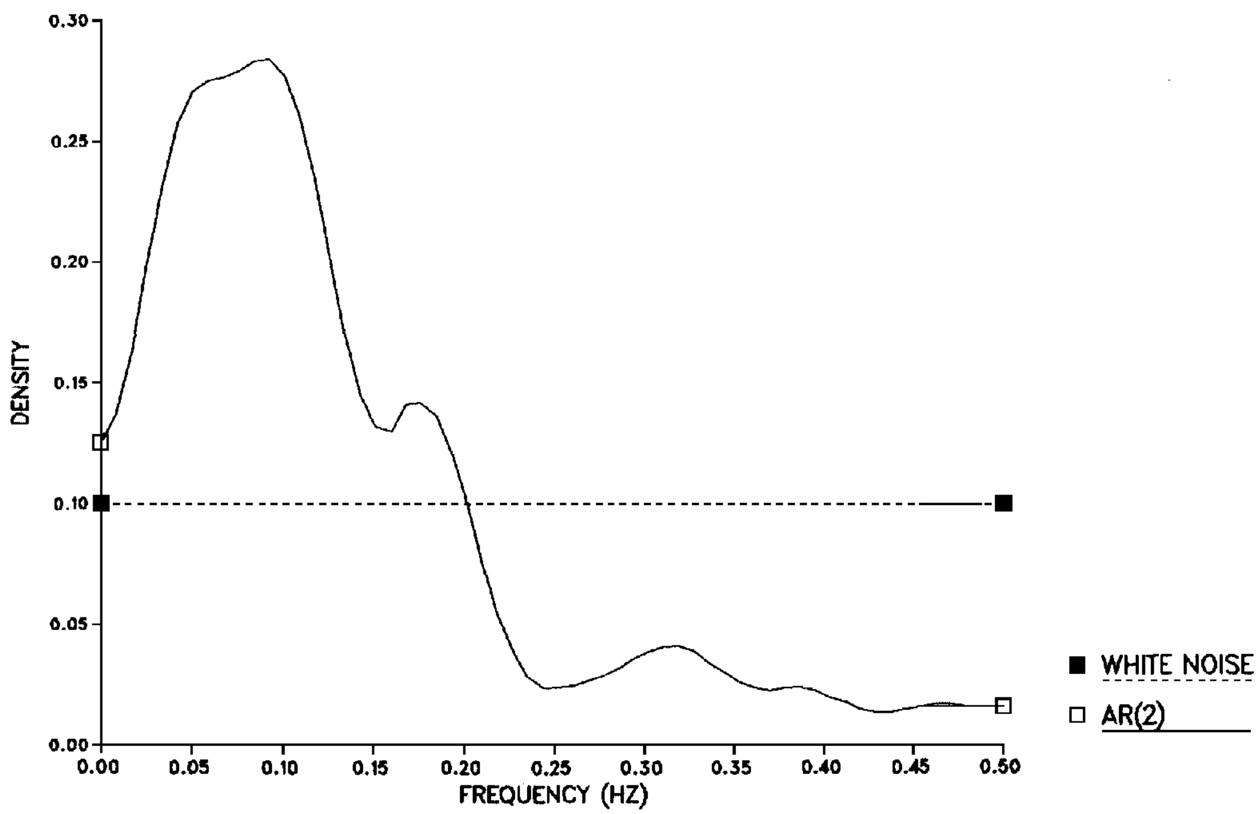

Figure A1. Spectral density function for a stochastic organization with one nonperiodic cycle.

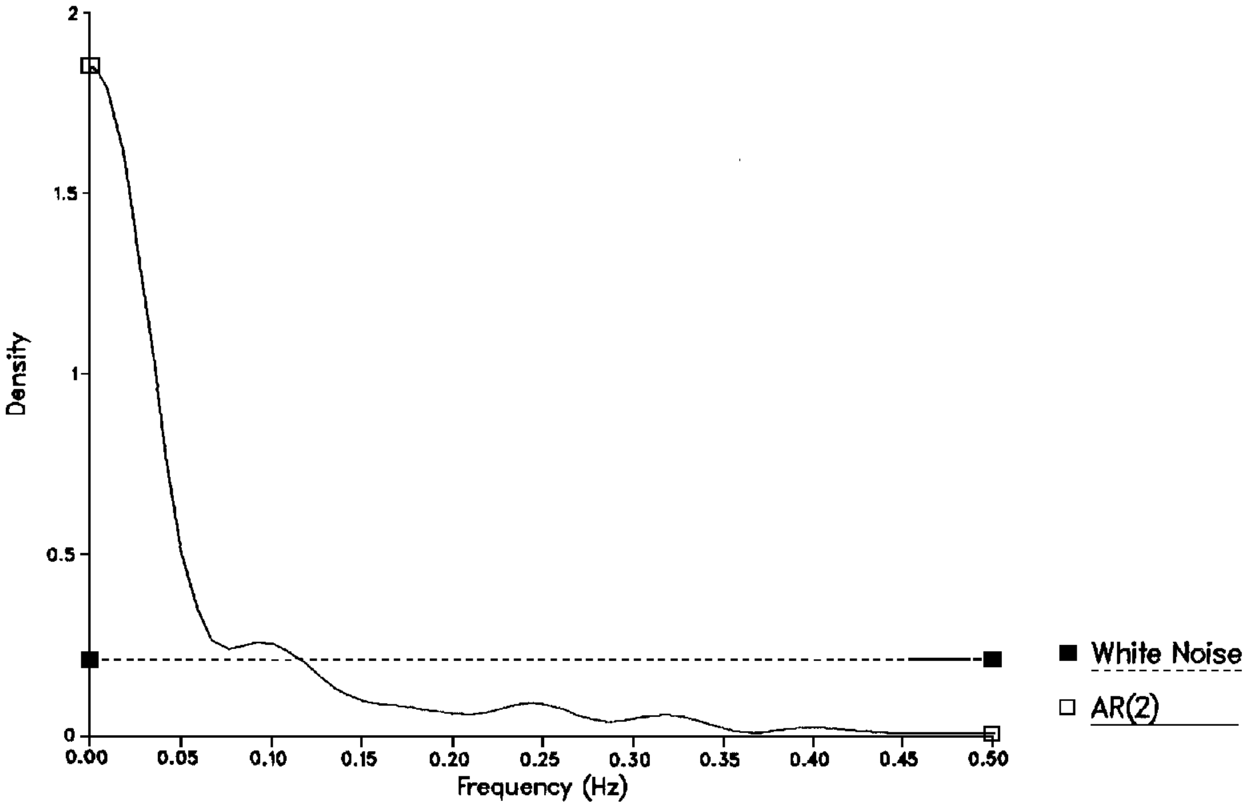

Figure A2. Spectral density function for a stochastic organization with no cycles.

When $a_1^2 + 4a_2 > 0$, an AR(2) organization is not cyclic, and the spectral density function will have no peak and be similar to that of an AR(1) organization, as exemplified in our data by Figure A2.

By contrast, the expected spectrum for a periodic cycle would peak over a narrow band of frequencies. (Were it not for measurement error, the spectrum for a periodic cycle would peak at a single frequency.)

Note that in both Figures A1 and A2 more than one frequency departs from the expectation for a random organization (i.e., white noise), as depicted by the dashed line in each figure. Significant spectral results cannot be interpreted as de facto evidence of periodicity. (See Gottman, 1981, or Chatfield, 1980 for further reference.)

Appendix B

An integrated process is one that is nonstationary in mean, variance, or covariance and must be differenced prior to making estimates of AR or MA parameters. A time series generated by both stochastic and periodic processes would have orthogonal AR or MA parameters for each component.

Autoregressive Integrated Moving Average (ARIMA) models are written in the form, (p, d, q) (p, d, q), where p refers to the order of the autoregressive parameter(s); d, to the order of differencing (typically $d = 0$ or 1); and q, to the moving average parameter(s). The repetition of terms is for periodic processes; when they are all zero, they are usually omitted from notation. A stochastic autoregressive process would be written in the form $(p, 0, 0)$ $(0, 0, 0)$, or more briefly as $(p, 0, 0)$ or $AR(p)$. A process with one autoregressive, stochastic component and one periodic component would be of the form $(p, 0, 0)$ $(p, 0, 0)$.

Univariate ARIMA modeling was performed using the time-series program in Minitab (Ryan, Joiner, & Ryan, 1982) as implemented on the University of Pittsburgh's DEC-10 computer. The program subtracts the mean from each observation prior to analysis.

Following McCleary and Hay (1980), we used the following univariate model-fitting procedure. First, each mother's and infant's series was examined for stationarity. A stationary series will have an autocorrelation function (ACF) that approaches zero as the distance between correlated observations (i.e., lag) increases (Gottman, 1981). Any series for which the ACF did not behave in this way was differenced before making parameter estimates.

Second, both the ACF and the partial autocorrelation function (PACF) for each series were examined. Different types of time-series have characteristic estimated ACFs and PACFs. For instance, the estimated ACF for a random process is zero; the estimated ACF for a nonperiodic AR(1) process decays exponentially. The estimated ACF for a periodic process will decay exponentially at periodic lags. For each series, after inspection of the ACF and PACF, a model was selected and fitted to the series. Model parameters had to be statistically significant, and the ACF of the residual series had to be acceptable by the following criteria: By visual inspection, no more than alpha (i.e., $\alpha = .05$) of the cross-correlations for the residual series could be significantly different from zero; cross-correlations for Lags 1–3 could not approach significance; and the Box-Pierce statistic, which is distributed as chi-square, had to be nonsignificant.

cance; and the Box-Pierce statistic, which is distributed as chi-square, had to be nonsignificant.

Appendix C

We used Gottman's (Williams & Gottman, 1982) spectral analysis program, which uses a Tukey-Hanning window. In general, we were not concerned with significance testing, because in each case the form of the spectral density function was known from the preceding time-domain analysis. When an unanticipated peak in the spectrum did occur, we calculated the period of the cycle in seconds and then refit the time-domain model with a periodic term at that lag. The parameter estimate for that lag was then tested for statistical significance ($p < .05$). An alternative test of significance would be to compute confidence intervals around the white-noise spectrum value.

Appendix D

Bivar is a bivariate, time-domain procedure that regresses each partner's behavior on both its own past (autoregression) and the past behavior of the other (cross-regression). It is an extension of the univariate modeling described in Appendix B with the following limitations: Differencing, if necessary to achieve stationarity, is assumed rather than included as a formal parameter; stochastic AR but not MA parameters are included; and no periodic parameters are included. The absence of MA parameters is of little consequence. The absence of periodic parameters would be, however, were a series strongly periodic.

Bivar uses a step-down procedure in which auto- and cross-regressive terms are systematically tested for significance and dropped as indicated. The bivariate models are: $M_t = \Sigma a_i \times M_{t-i} + \Sigma b_i \times B_{t-i} + e_t$, and $B_t = \Sigma c_i \times B_{t-i} + \Sigma d_i \times M_{t-i} + n_t$, where B_{t-i} = Baby's behavior at Time $t-i$, M_{t-i} = Mother's behavior at Time $t-i$, and e_t and n_t = random noise or error at Time t. The baby's (or mother's) behavior is a function of both her own previous behavior and that of her partner. For both the mother and the infant, a likelihood-ratio procedure tests the significance ($p < .05$) of the difference between the larger bivariate and the smaller univariate model. The bivariate models reduce to univariate models when the cross-regressive coefficient is not significantly different from zero.

To reduce the likelihood of spurious findings, the univariate modeling and inspection of the CCF guided our choice of the number of parameters at which to begin the step-down procedure. The initial number of AR terms was set at one more than that found in the largest univariate AR model, which was a third-order autoregressive model—AR(3). The number of cross-regressive terms was individually set according to the CCF for a given dyad. For instance, if the CCF suggested that the baby was following the mother at a lag of 1–6 s, we set the initial number of cross-regressive terms at *6*.

Bivar does not test for simultaneous cross-lag dependence. Most other bivariate time-series programs do (e.g., BMDP 2T), however. To correct this omission, we modified Bivar accordingly. We also subtracted the mean from each observation prior to analysis to aid interpretation and for consistency with the univariate analyses.

Received December 8, 1986
Revision received October 2, 1987
Accepted November 25, 1987 ■