Interaction Diffusion and Fertility Transition in Costa Rica*

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Abstract

A long-standing concern of sociologists is the contribution of diffusion processes to social change. This article considers the contribution of social interaction diffusion to the fertility transition in Costa Rica, focusing on person-to-person contagion. Several prominent features of the Costa Rican transition suggest the existence of interaction diffusion effects, notably its pervasiveness toward all socioeconomic strata and the lack of evidence of a downward shift in family size preferences. Maps of the timing of fertility transition show an ordered spatial pattern suggestive of contagion between neighboring areas. A dynamic regression model estimated from pooled time series data for 100 counties reveals inter- and within-county diffusion effects on birth control adoption net of socioeconomic and family-planning program effects.

This article tests the hypothesis that diffusion of information and values through social interaction shaped the timing and the pace of the transition from high to low levels of fertility in Costa Rica. Previous studies of Costa Rica clearly document that the adoption of birth control behavior occurred earlier in some segments of the society than in others (Behm & Guzmán 1979; Stycos 1982), and hence it is not controversial to apply the term *diffusion* in a descriptive sense to the Costa Rican transition. The aim of this article is to test a more demanding proposition, namely that diffusion dynamics explain aspects of the transition that *cannot* be explained by changes in the socioeconomic structure of the society nor by changes in the availability of birth control. By invoking plausible and straightforward behavioral models, we posit that birth control can be "contagious" under certain conditions. If so, diffusion dynamics constitute a further type of causal agent of fertility transition, distinct from decreased demand for children due to structural changes and increased supply of birth

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control methods. This hypothesis is tested by estimating dynamic regression models on a pooled time series for 100 areas (a combination of "cantones" and "distritos") of Costa Rica for the period from 1958 to 1988.

The hypothesis that diffusion dynamics partially account for the timing and pace of the decline of fertility in Costa Rica is motivated by fundamental features of the decline that are reviewed in the next section of this article. Similar hypotheses have emerged as a common theme in research on fertility transition during the past decade and, furthermore, can be grounded in several major streams in sociological thought.

The considerable attention given to diffusion processes in the recent literature on fertility change can be attributed to several influential interpretations of the findings of the Princeton European Fertility Project (van de Walle & Knodel 1979; Watkins 1987). Specific features of the fertility decline in Europe suggest that it cannot be simply explained by societal structural changes but rather that it was shaped by diffusion processes. Among the key pieces of evidence cited are the apparently autonomous spread of knowledge and practice of birth control, the pervasiveness of this process across socioeconomic groups, the strong patterning of decline in terms of cultural and linguistic boundaries, and the character of the transition from high to low fertility as a "one-time event." Following similar lines of argument, Cleland and Wilson (1987) argue that the results from the mammoth World Fertility Survey are incompatible with theories of fertility transition that stress the primary causal role of socioeconomic changes that alter the demand for children. This line of reasoning can lead to a counterproductive opposing of socioeconomic and diffusionist explanations. The most influential example of this dichotomization is Carlsson's 1966 article that has structured much of the subsequent debate. Carlsson distinguishes between "innovation" and "adjustment" explanations for the widespread adoption of birth control practices. Under the adjustment explanation, adoption of birth control is a response to changed socioeconomic conditions that makes use of already existing behavioral choices. Carlsson's concern is whether or not birth control is innovative behavior, not the process of diffusion of that innovation. A related literature proposes that ideational change rather than change in material circumstances is the driving force behind fertility transition (Cleland & Wilson 1987), with the diffusion of ideas thus assuming a critical causal role. Freedman (1987) also attributes an important causal role to ideas and mental frameworks that spread through national and world networks of communication and transportation, although he is less dismissive of the contribution of changes in material circumstances than are Cleland and Wilson. At the extreme, some literature subsumes all explanations for fertility change within a diffusion paradigm (Rogers 1973, 1983). Individual acceptance or rejection of family planning is explained not in terms of motivational forces or socioeconomic conditions, but in terms of the "adoptive potential" of the innovation (i.e., as an essential attribute of birth control) and individuals' "innovativeness." Taken to this extreme, diffusion is more a way of thinking than a category of analysis.

Reading the recent demographic literature one would be forced to conclude that it has emerged largely in isolation from major streams in sociological thought that have grappled with some of the same basic questions about the causes of social change and have offered explanations similar in form and content. Diffusionist models of social change have a long history in sociology (see a concise review in Smith 1976). The fundamental premise of these models is that social change occurs, in part at least, as a result of exposure of one social entity to the values, behaviors, and technologies of another. Often the epidemiological term *contagion* is invoked, and occasionally the term *imitation* is also employed (e.g., Tarde 1890). While these models have been largely out of favor in the postwar period among macro theorists, elements from them can still be found in works such as Bendix's *Kings or People* (1978). Coleman (1964) pioneered in introducing sociologists to the formal mathematics of diffusion processes (Hamblin, Jacobsen & Miller 1973; Tilly 1984). Variants of the models sketched by Coleman have been applied in empirical research on collective violence (Pitcher, Hamblin & Miller 1978), government reform (Knoke 1982) and participation in formal religious organizations (Land, Deane & Blau 1991), to cite three examples. The analysis of the Costa Rican fertility decline presented in this article uses models of the same general form.

Most of the macroresearch on diffusion processes presumes, although it is often not acknowledged, social influence processes operating at the microlevel. The massive social psychological literature on social influence therefore provides the microfoundations for the macrodiffusion models. French's (1956) seminal work articulates the issues that continue to be addressed in recent research on social influence (e.g., Friedkin & Cook 1990). Bandura (1986) and Moscovici (1985) provided more recent summaries of these research traditions. Watkins (1993) reviews the more limited evidence on social interaction effects on fertility change in contemporary developing countries. She concludes that there is empirical support for the view that social interaction processes affect fertility behavior. For our research on fertility transition, the important contribution of this social psychological literature is that it specifies the behavioral mechanisms through which the spatial-temporal patterns that we observe in Costa Rica might be generated. This literature also underscores the importance of moving beyond aggregate-level modeling to direct investigation of the exertion of social influence effects through local social networks, a conclusion that Entwisle et al. (1992) also arrive at through analysis of unstructured interviews on contraceptive behavior in rural Thailand. Our future research will move in this direction. For now, we note only that the extensive sociological literature on social influence processes confirms the plausibility of the behavioral dynamics that we hypothesize as having affected the timing and pace of the Costa Rican fertility transition. We are not able to adjudicate among contending hypotheses about the precise mechanisms underlying social influence effects. This problem has been the focus of recent microlevel research (e.g., Burt 1987; Friedkin & Cook 1990).

It is also worth noting that debate about the magnitude of social influence effects has occupied a prominent place in the sociological literature on socioeconomic status for at least three decades. It may be recalled that strong claims concerning peer-group influence were among the more controversial elements of the 1966 Coleman report on educational achievement (Coleman et al. 1966). Empirical research on this topic has continued up to the present (e.g., Bank et al. 1990; Cohen 1983; Davies & Kandel 1981). A related literature that has emerged more recently examines the magnitude of "neighborhood effects" on status attainment and adolescent fertility, among other variables (e.g., Case & Katz 1991; Evans, Oates & Schwab 1992; Jencks & Mayer 1989). The intensity of the debate in this literature about social influence effects on young adult achievement reflects the high public policy stakes. For similar reasons, the debate in the demographic literature about diffusion effects on fertility transition has at times become quite vociferous.

We turn now to the conceptualization of diffusion effects on fertility transition. The central hypothesis of this article is that social influence effects (social contagion) shape fertility transition. More precisely, the key proposition is that the adoption of birth control by some individuals influences the likelihood of adoption by others. Elsewhere such a social dynamic has been termed "endogenous feedback" (Erbring & Young 1979). Three behavioral mechanisms might account for such a feedback process (Casterline & Knight 1993): (1) *Information flow:* when individuals are uninformed about available birth control technologies, those who learn about, or who adopt, certain methods can serve as sources of information for others; (2) *Demonstration effect:* when individuals are uncertain about the benefits and costs of certain fertility decisions, the experiences of others provide concrete evidence of the benefits and costs of the same choices; (3) *Changes in normative context:* when social norms prohibit certain reproductive behaviors such as the use of modern contraceptives to limit family size, individual decisions to adopt innovative behavior can modify the group norms. Whatever the relative contributions of these behavioral mechanisms, the main thrust of the diffusionist argument as formulated here is that individuals are not isolated, that they talk to each other and observe each other, and in doing so transmit information, beliefs, and values on reproductive matters.

We focus on a specific type of diffusion, namely that generated internally in a system by social interaction. Other types of diffusion processes occur from external sources (e.g., the mass media), marketing efforts of professional propagators, and the institutional spread of propagator agencies (Brown 1981; Coleman 1964; Mahajan & Peterson 1985). In analogy to epidemiology, this article is concerned with the spread of diseases by person-to-person contagion rather than by exposure to fixed hazards such as a contaminated well. This focus on person-to-person spread does not deny the importance of other channels of diffusion such as the mass media. However, external-influence diffusion can be accommodated within conventional explanatory frameworks for fertility (e.g., Easterlin's synthesis framework) as another set of factors influencing the costs and benefits of children and/or the psychosocial or material costs of contraception. Social interaction diffusion, by contrast, adds a further causal dimension: adoption of contraception depends not only on personal and institutional factors but also on the extent to which other individuals have already adopted birth control and the degree of social interaction between adopters and nonadopters. An earlier article shows through formal modeling that social interaction diffusion has the potential to powerfully affect the timing and pace of fertility transition (Rosero-Bixby & Casterline 1993).

Fertility Transition in Costa Rica

Costa Rica experienced one of the earliest and fastest, although incomplete, fertility transitions in the developing world. The total fertility rate (TFR) fell from 7.3 to 5.5 between 1960 and 1968, the year when an energetic national family planning program started, and then to 3.7 in 1976, the year when the decline abruptly stopped (United Nations 1985). Only the fertility shifts in Singapore and Taiwan were faster, in a similar period, than this reduction of 3.6 births between 1960 and 1976 (Coale 1983). During the decade following 1976, however, the total fertility rate fluctuated erratically around 3.7 births, and the contraceptive prevalence rate oscillated around 65% to 70% (Asociación Demográfica Costarricense 1987). A moderate declining trend was in place again in the late 1980s, reducing the total fertility rate to 3.2 in 1990, according to the National Directorate of Statistics and Census. The following review sorts out five elements in the Costa Rican fertility transition that, taken together, suggest diffusion dynamics: (1) absence of birth control before the fertility transition; (2) limited change in family size preferences during the transition; (3) rapid pace of fertility change relative to the pace of socioeconomic development; (4) pervasiveness of the decline; and (5) ordered spatial pattern of decline.

PRETRANSITION FERTILITY

Determining whether birth control was known and accepted before the transition is central in the debate about whether fertility change reflects the diffusion of innovative birth control behaviors or rather reflects greater prevalence of preexisting behaviors in response to changes in societal structures (Carlsson 1966; Knodel 1977). Retrospective estimates show that by 1960, the approximate year when the transition started, the total marital fertility rate was about 8 births, the average length of breast-feeding was less than 6 months, and the contraceptive prevalence rate was 16% (Rosero-Bixby & Oberle 1989). This high level of marital fertility is not compatible with generalized birth control practices. The relatively brief average duration of breast-feeding is in part responsible for this high marital fertility. The contraceptive prevalence rate of 16%, although low, suggests the existence of strata in which birth control practice was common. These strata were emerging urban middle classes, using almost exclusively condoms and the rhythm and withdrawal methods (Rosero-Bixby, Gómez & Rodríguez 1982). It is not clear why birth control did not diffuse from these strata to other strata before 1960. Some suggest that the event triggering the diffusion process was the commercial introduction of oral contraceptives in 1962 and the subsequent media attention given to the new contraceptives (TinMyaingThein & Reynolds 1972).

FAMILY SIZE PREFERENCES

Surveys carried out during the course of the fertility transition show little change in family size preferences. In the first fertility survey conducted in Costa Rica in urban areas in 1964, women younger than 35 stated a desired family size of 3.9 children on average. The corresponding average for the first rural survey

(1969) was 4.6 children. Retrospective cohort-specific estimates indicate that the average desired family size for women aged 20 to 24 in 1954 was 3.9 children in urban areas and 4.5 children in rural areas (Rosero-Bixby, Gómez & Rodríguez 1982: Tables 5.3 and 5.8), i.e., similar to those desired sizes for 1964 and 1969. As of 1986 these averages have declined by less than one child, to 3.1 and 4.1 children in urban and rural areas, respectively (Hermalin, Riley & Rosero-Bixby 1989: Table 1). Thus, there is no evidence of a downward shift in family size preferences in the decade prior to the fertility transition, nor is the substantial decline in fertility accompanied by a decline in preferences of corresponding magnitude. The contrast between the modest change in family size preferences and the sharp decline in fertility suggests that motivational theories alone have only a limited explanatory power in Costa Rica during the period of sharpest fertility decline.

Moving further back in the past, in a 1944 book two U.S. sociologists report that well-educated Costa Ricans had a clear preference for small families: "the great majority of young couples prefer to have not less than two or more than four children . . . Of 100 university students, 72 say large families should be avoided; the boys are more strongly (94%) in favor of limitation. The average student wants three children" (Biesanz & Biesanz 1944:74-75). Although this book provides less clear evidence about the fertility preferences of the working classes, the following quotation from a 60-year-old woman is suggestive of a latent demand for family planning:

> After I had had three children I was thin and sick and didn't want any more. A friend came to me and said, "Why don't you tell your husband to do thus and so?" I did, and he was furious. "You have to be a *good* woman. Don't ever see that woman again." I prayed not to have any more and tied a cord of San Francisco tight around my waist. But I had twelve more. Birth control is a sin. Everybody does it now but they don't confess. (Biesanz & Biesanz 1944:75)

Note that for this woman the obstacle to birth control does not appear to be knowledge of contraceptive techniques but rather social norms, enforced by her husband, about the unacceptability of birth control. We posited above that social interaction diffusion dynamics can modify norms about the acceptability of contraception.

It thus seems that for many years Costa Rican women wanted families of moderate size but had large families. Early surveys show 47% of urban respondents (1964) and 78% of rural respondents (1969) expressed a desire to regulate their fertility but did not use contraception (Gómez 1989: Table 4.7). These figures leave no doubt about the existence of a substantial discrepancy between stated fertility desires and behavior during the early years of the Costa Rican fertility transition, especially in rural areas. However, by 1976 the comparable figures were 32% in urban areas and 38% in rural areas, and by 1986, 20% and 27%, respectively. This sharp reduction in the discrepancy between fertility preferences and behavior suggests that factors other than motivation made a major contribution to the Costa Rican transition. Possible factors include an increased availability and acceptability of contraceptives and the interaction diffusion effects postulated in this article.

RELATIVE SPEED OF THE DECLINE

Fertility in Costa Rica fell by 50% in only 15 years. In those years the country's economy grew substantially. Real growth in the gross domestic product was 6.5% per annum in the 1960s and 1970s (Lundahl 1991). The pace of social development was even faster than that of economic growth. Stycos (1982) singles out increased schooling as a principal cause of the fertility decline: "between 1950 and 1963, and again between 1963 and 1973, there were huge increases in the proportion of young women completing primary education: from 22% in 1950, the proportion rose to 40% in 1963, and to 72% in 1973 . . . by 1963 both the pervasiveness of literacy and the substantial increases in the proportion completing elementary school might, indeed, have signaled the fertility decline of the 1960s" (17-18). These rapid increases in income and educational attainment undoubtedly stimulated the spread of birth control behavior in Costa Rica. But several efforts to quantify their impact suggest that they explain only a small part of the fertility decline. One study using census data concludes that 22% of the fertility decline between 1960 and 1970 is attributable to schooling improvements (Behm & Guzmán 1979). Another study using survey data concludes that changes in educational attainment, occupational structure, and women's employment together explain 22% of the fertility decline in urban areas from 1961 to 1973 and 15% of the decline in rural areas from 1966 to 1973 (Rosero-Bixby et al. 1982:41).

PERVASIVENESS ACROSS SOCIOECONOMIC STRATA

The failure of socioeconomic changes to explain a large fraction of the fertility decline simply reflects the fact that all socioeconomic strata participated in the decline. In the early 1960s, socioeconomic differentials in fertility were substantial and widening (Behm & Guzmán 1979). Fertility decline began, and initially was limited to, the urban middle classes, the urban educated in particular. In the 1970s the process reached all strata. Even among wives of illiterate farmers, a majority were adopting contraception by the end of the 1970s (Rosero-Bixby et al. 1982; Stycos 1982). The rapid spread of birth control practices to all socioeconomic strata is one of the most striking features of the Costa Rican fertility transition. A pattern in which innovative behavior first adopted by an elite *rapidly* pervades all strata is strongly suggestive of a causal role of social contagion. Were the timing and pace of the decline determined primarily by the socioeconomic structural changes also underway in this period, one would expect birth control adoption to be explained mostly by the shrinking of lower, more traditional strata.

SPATIAL SPREAD

The evolution of geographic differentials in fertility was analogous to the evolution of socioeconomic differentials: they increased at the beginning, reaching a peak as of 1966-67, and then narrowed (Klijzing & Taylor 1982). A study of these spatial patterns states that "adoption of birth control occurred first in the nation's main innovation center, San Jose, and gradually spread toward the periphery" (Fridman 1984:169). The pattern of spread of birth control practices to all areas of the country is once again suggestive of the contribution of interaction diffusion dynamics.

Analytical Model for Testing Diffusion Effects

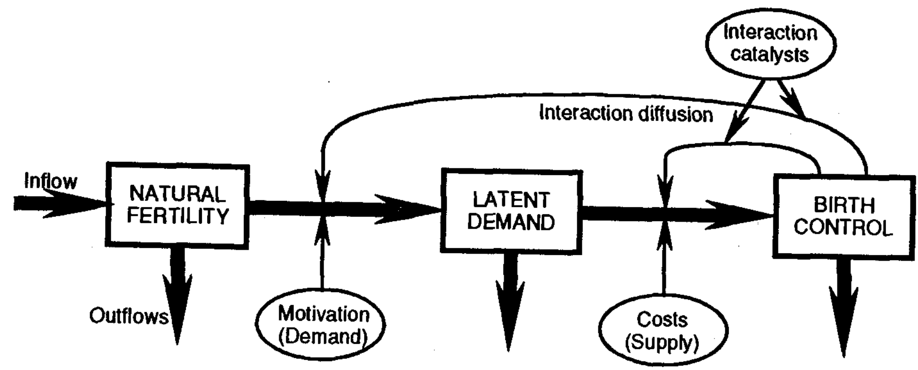
Figure 1 shows a simplified model for fertility. In order to portray the diffusion feedback, the diagram presents fertility as a dynamic system. There are three states: natural fertility, latent demand (i.e., motivation to regulate fertility but no birth control practice), and birth control. From the population perspective, couples flow into and out of this model, as they age into and out of the reproductive years. Within the model, couples can first move from the state of natural fertility to latent demand as they develop a motivation to restrict their fertility, and from the state of latent demand to birth control. In pretransition societies the flow into the birth control state is null (or almost null), and the whole population exists in a regime characterized by an absence of deliberate fertility limitation. In such populations, the fraction of couples in the natural fertility state as against the latent demand state is of some significance for the potential for substantial (and rapid) fertility decline, but this breakdown has no impact on the observed patterns of fertility. In posttransition populations, most couples remain in the birth control state, into which they probably move early in their reproductive lives.

Fertility demand and birth control supply, as well as a diffusion feedback, drive the flow among the three states. Socioeconomic and cultural forces determine demand. The supply forces, in turn, comprise cost constraints to the use of contraception, including normative, informational, psychological, and objective barriers; these are the costs that family planning programs strive to reduce. The diagram portrays diffusion as a feedback emanating from the individuals in the birth control state, conforming to the hypothesis stated above that the adoption of birth control by some individuals influences the likelihood of adoption by others. The more couples in the birth control state the stronger is the feedback effect. The population size in the natural fertility state also determines the size of the flow, which must be small when most couples have already moved to the birth control state. The diagram also specifies that some factors act as catalysts of diffusion; these are determinants of the intensity of social interaction, such as geographic compactness, social homogeneity, and the taboo character of some reproductive topics, as postulated by Retherford and Palmore (1983). This model is described more fully, and its formal implications developed in detail, in Rosero-Bixby & Casterline (1993).

The system in Figure 1 is for a closed population with homogeneous mixing, which means that an individual is equally likely to interact with any other member of the population. A more realistic representation must include diffusion effects from the outside world, i.e., from those in the birth control state in other populations. In doing so, the homogeneity assumption is implicitly removed. In an open system, the degree of openness of the population will condition the effect of out-community contagion.

A short-hand mathematical representation of the dynamic model in Figure 1 is as follows:

FIGURE 1: A Causal Diagram for Birth Control Adoption



$$\frac{\partial Y}{\partial t} = f(X, Z, Y, Y^2, \ddot{Y}, H) \qquad (1)$$

where the first derivative $\partial Y/\partial t$ represents the increase per unit of time in the number (or in the cumulative proportion) of adopters of birth control, i.e., the flow into the birth control state. This flow is a function of a vector X of socioeconomic and cultural variables (the demand factors) and a vector Z of costs of birth control (the supply factors). It is also a quadratic function of the cumulative number of adopters Y in the community; as well as a function of the cumulative number of adopters $\ddot{Y}$ in other relevant communities. In addition, a vector H of diffusion catalysts, such as the openness and the social and geographic compactness of communities, condition the rate of diffusion.

Assuming linearity and two-year lags for the demand and supply effects, and expressing the model in its discrete, recursive, and stochastic formulation, one arrives at the following equation for testing interaction diffusion effects:

$$Y_t = \kappa + \alpha X_{t-2} + \beta Z_{t-2} - \lambda_1 Y_{t-1} + \lambda_2 Y^2_{t-1} + \lambda_3 \ddot{Y}_{t-1} + \lambda_4 H_t Y_{t-1} + \lambda_5 H_t \ddot{Y}_{t-1} + U_t \qquad (2)$$

The Greek letters are the parameters to estimate and U_t represents the error term. The parameters λ denote diffusion effects. For convenience, the subscript i for county is omitted. Testing the proposition that social interaction influences the adoption of birth control is thus equivalent to testing whether the vector of parameters λ is different from 0. With the exception of λ_2 (see below), all λ are expected to be positive. Note that equation 2 allows for the interaction diffusion effects to be conditioned by the catalysts H_t.

The inclusion of lagged values of the dependent variable Y_{t-1} in the right-hand side of the equation makes the model dynamic (Hanushek & Jackson 1977). The effects of the X and Z variables are net of the level of birth control during the previous year (Y_{t-1}), and hence for these two blocks of variables this is a model not for cross-sectional variation in Y but rather for its variation over time.

The coefficients for the lagged value of Y and Y^2 represent the endogenous feedback, or within-community, diffusion effect. Inclusion of the quadratic term allows for a decaying contagion effect, reflecting a shrinking pool of potential adopters as birth control practice increases. Were the quadratic term to be omitted, the equation would suggest external-source diffusion only (Huckfeldt, Kohfeld & Likens 1982; Mahajan & Peterson 1985), which is not of interest here. The presence of a quadratic term makes the within-community diffusion effect conditional on the level of birth control. This effect is given by the derivative:

$$\frac{\partial Y_t}{\partial Y_{t-1}} = \lambda_1 + \lambda_4 H_t + 2\lambda_2 Y_{t-1}$$

If contagion dynamics are present, λ_1 should be positive in sign and λ_2 should be negative, i.e., the feedback effect declines over time owing to a shrinking pool of couples in the natural fertility state.

Estimation of the Model: The Data

The data for estimating equation 2 consist of pooled time series for 100 counties during the period 1958-88. To maintain two-year lags for X and Z, the first two years of Y_t must be omitted, resulting in a final data set consisting of 2,800 observations (28 years x 100 counties). Counties were defined specifically for this analysis. They combine the formal administrative units of Costa Rica termed *canton* (81 units) and *district* (421 units), following in part the work of the German geographer Nuhn (1978), who proposes a division of Costa Rica into 85 small regions, with approximate canton boundaries. Two principal criteria govern the definition of the counties: first, maintenance of the same boundary layout during the entire period; second, having units that are geographically homogeneous, well integrated, and thus socially meaningful. The population size of the counties generally ranges between 5,000 and 50,000; their territory averages about 100 square km in the highlands of the Central Valley and 900 square km in the lowlands. (For further details on the county scheme, see Rosero-Bixby 1991.)

No data are available on birth control use at the county level. However, using vital statistics and census data, general marital fertility rates (births per women married or in consensual unions, aged 15 to 44) can be calculated at the county level for the years encompassing the fertility decline. These rates can be used as a proxy for the extent of birth control practices. To assess the validity of these rates, we compared them against county-specific estimates generated by retrospective projections of births from the population under five years of age in the 1963, 1973, and 1984 censuses. This comparison revealed that in three counties — two of them in the capital city — there was a probable over-registration of births, whereas in 15 counties there was a probable under-registration. The time series for these 18 counties were corrected accordingly (Rosero-Bixby 1991).

Comparison of the national time series of marital fertility rates and contraceptive prevalence rates shows a near perfect correlation between the two series (Rosero-Bixby & Oberle 1989). A second comparison of time series of the general marital fertility rate and the Coale index of marital fertility (Coale 1967) for the seven provinces of Costa Rica also shows a near perfect correlation. These comparisons suggest that the fertility rate series used in this analysis are a good proxy for the prevalence of birth control, in particular that they do not appear to be distorted by shifts in nuptiality patterns or age composition. To facilitate the interpretation of results, the marital fertility rates (F) are transformed into the birth-control variables Y and $\ddot{Y}$ assuming a simple linear relationship, which approximates the relationship observed for the national time series of marital fertility rates and contraceptive prevalence rates:

$$Y = 1.25 - 2.5F; \qquad Y \geq 0$$

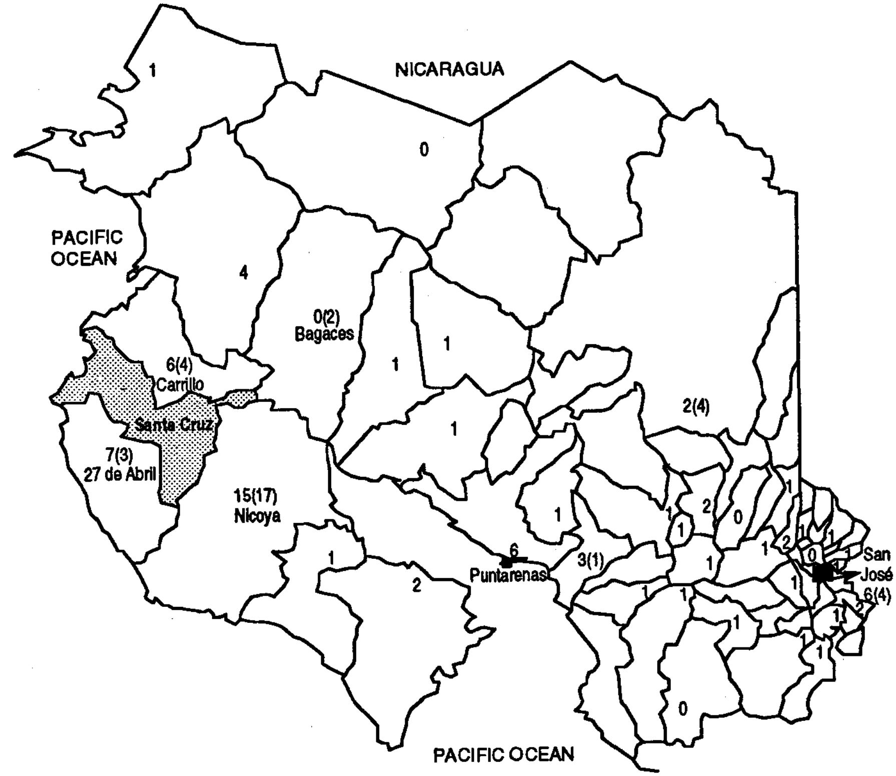
To meet the condition $Y \geq 0$, all fertility rates 0.5 or larger were transformed into $Y = 0$. Hence variations at very high levels of fertility (above 0.5) are not permitted to affect the analysis, under the assumption that they reflect behavior other than birth control practices. Since fertility rates lower than 0.2 were rarely observed, the upper limit of the proportion Y is about 0.8. Because Y is just a linear transformation of F, analyses based on the estimated proportion Y will produce similar results as analyses based on fertility rates. The parameters of an equation for Y are easier to interpret in terms of the diffusion of fertility control.

In addition to the one-year lagged Y and Y^2, two other sets of variables are used in the testing for diffusion effects.

OUT-COUNTY BIRTH CONTROL

Out-county birth control Y, which is measured as the weighted average of Y in all other relevant counties, with weights (summing to 1.0) proportional to the amount of social interaction of each county with the index county. The weighting matrix draws on the concept of *mean information field* (MIF) developed by geographers for the study of spatial diffusion (Hägerstrand [1953] 1967). The MIF concept comes from the observation that individuals have a spatially biased field of information, which decays with the distance from their home. Geographers have generated MIF estimates from such data as telephone calls, local migration, and daily travel patterns (Marble & Nystuen 1963.) For this study, a matrix of social interactions among the 100 counties was estimated with 1984 census data on the birth place of spouses (marriages or consensual unions) in the age bracket 20 to 34. The estimated interactions between the 4,900 pairs of counties were validated with a gravity model (Haynes & Fotheringham 1984). Gravity models postulate that interactions between two places increase with their demographic "masses" and decay with the distance between them. The goodness-of-fit of a simple gravity model to our interaction estimates is equivalent to a multiple correlation of 0.83, indicating that the matrix of social interactions conforms roughly to the predictions of the gravity model. This can be interpreted as a sign of overall reliability. To illustrate, the map in Figure 2 shows the interaction weights for Santa Cruz county derived from marriage distances and, in parentheses, those predicted by the gravity model. About 40% of all Santa Cruz interactions occur within just five counties, three of which border on Santa Cruz: Nicoya (15%), 27 de Abril (7%) and Carrillo (6%). The other two are located farther away but have a large demographic mass: Puntarenas, the most important Pacific port (6%), and San José, the capital city

FIGURE 2: Map of Interaction Weights for Santa Cruz County[a]



[a] The numbers indicate marriage interactions (per 100). The expected interactions according to a gravity model are shown in parentheses, if these differ from the observed interactions by more than one percent point.

(6%). The remaining 60% of the interactions are dispersed among 57 counties (some of them not shown on the map).

DIFFUSION CATALYSTS H

We define three indicators (1) *Openness*, represented by the proportion of natives who marry a spouse from other counties; this is calculated from the data matrix described in the previous paragraph. (2) *Social compactness*, measured by the homogeneity (one minus the coefficient of variation) in the number of years of educational attainment of adults born between 1923 and 1943, i.e., in their prime reproductive ages in the early 1960s. (3) *Geographic compactness*, calculated as one minus the standard deviation in the travel time to San José across census tracts in the county (using unpublished tract-level data from the 1984 census). Note that all three catalyst variables are measured for one historical period only, i.e., do not vary over time, contrary to the specification of equation 2. We are not able to locate time-varying H variables, which is regrettable given the conclusions in our previous article (Rosero-Bixby & Casterline 1993) that powerful social interaction diffusion effects probably reflect intensified contact among societal subgroups, a change that in turn can be attributed in part to changes in the economy and in basic infrastructures (e.g., transportation and communication systems).

The catalyst variables H enter equation 2 only as conditioning influences on within-county and across-county diffusion effects, i.e., in interaction with Y_{t-1} and $\check{Y}_{t-1}$. The diffusion hypotheses are that within-county contagion is stronger in counties that are geographically and socially compact, and that across-county contagion is stronger for counties that are more open to the outside world.

The vector of demand variables X consists of time series of four indicators derived from the 1963, 1973 and 1984 censuses. We do not comment on hypothesized effects here, as these variables are not the focus of this analysis. The four indicators are socioeconomic development, child mortality, social security coverage, and legal unions.

SOCIOECONOMIC DEVELOPMENT

An index constructed as a linear combination of seven items: household wealth, nonagricultural labor force, urbanization, labor force participation of women, literacy, school attendance, and educational attainment. This index was defined to avoid multicollinearity, after a principal component analysis showed considerable overlap among the seven variables. The index explains 85% of the joint variance of these seven variables. In contrast, the following three variables showed some degree of orthogonality to this index of development.

CHILD MORTALITY

The probability of death before age five. The series of yearly rates were interpolated from 10 points retrospectively estimated from the 1973 and 1984 census using the MORTPAK computer program (United Nations 1988).

SOCIAL SECURITY COVERAGE

Proportion of the population covered by the social security system, which provides both retirement benefits and health insurance.

LEGAL UNIONS

Proportion of women in a union who are legally (as opposed to consensually) married.

Finally, the supply variables Z are represented by a single indicator of potential physical *access to public family planning services*. This indicator takes into account the supply of services in a radius of 20 miles (weighted by the inverse of distance) and the size of the catchment population of each outlet, according to a formula proposed by Joseph & Bantock (1982). The indicator is calculated from information on the size of clinics and the year when family planning services were first provided (obtained from administrative files) and the geographic coordinates of the clinics and the census population. We discarded a simpler indicator of density of services in the county because it was clear that many people overrun counties' boundaries to obtain family planning service (e.g., there are many counties with no services, as well as many with a high density of services). For the same reason, and because of wild year-to-year fluctuations that suggest substantial reporting deficiencies, we discarded indicators based on the number of new adopters and consultations by clinic, which have been used by early authors (Fridman 1984; Stycos 1982).

The X and Z variables are lagged two years to allow for behavioral and biological lags in their effects (Schultz 1973). We resist more precise temporal specification through distributed-lag formulations, given that these variables have the limited function of controlling for forces that may confound or distort the estimation of diffusion effects.

In addition to the X and Z variables, we include a measure of the level of *natural fertility* and a *historical time*. The marital fertility rate of the county in 1958-62 proxies its natural fertility. To the extent counties vary in natural fertility, the estimation of county variation in birth control through a simple linear transformation of marital fertility rates is biased. Adjustment for preexisting levels of fertility should largely eliminate this bias. A negative regression coefficient is expected for this control variable. Historical time is entered as a single continuous term, therefore capturing national linear time trends in fertility and its determinants that reflect the impact of unmeasured variables that affect the country as a whole (such as the national mass media). Inclusion of this term provides a more stringent test of the net effects of the diffusion terms (Montgomery & Casterline 1993). To the extent that these developments lead to similar county-specific trends in marital fertility and in the X and Z variables, estimates of diffusion effects in equation 2 will be distorted.

Estimation Approach

Equation 2 expresses interaction diffusion effects in a conceptually straightforward fashion. Estimation of this equation poses serious challenges, however. Even without the autoregressive specification, ordinary least squares (OLS) assumptions would probably not apply, since the error term is likely to be characterized by within-county correlation over time (autocorrelation) and by nonconstant variance from one cross-section to another (heteroskedasticity) (Judge et al. 1985). In their simple forms these two problems can be addressed by adopting a "cross-sectionally heteroskedastic and timewise autoregressive model" (Kmenta 1986:618-22). This model can be estimated by making two successive transformations of the variables, one to purge the autocorrelation within counties and the other to take account of heteroskedasticity across counties. The transformations for all variables m in county i are:

(1) $m_{it}* = m_{it} - r_i m_{it-1}$, and

(2) $m_{it}** = \dfrac{m_{it}*}{s_{ui}}$

where r_i is the serial correlation coefficient for the residuals in each county, and s_{ui} is the standard deviation of the residuals U_{it} specific to each county. Ignoring for the time being the further complications introduced by the presence of lagged Y and by the possibility of error correlation across counties, one could consistently and efficiently estimate equation 1 by, first, obtaining r_i as the correlation coefficient between OLS residuals in consecutive years and, second, obtaining s_{ui} as the standard deviation of the OLS residuals after transformation 1. OLS estimation can then be applied to the doubly transformed variables. This is, in effect, estimation by generalized least squares (GLS).

Unfortunately, this approach probably does not yield consistent estimates of the parameters λ, for two reasons: first, the presence of the lagged endogenous variables Y_{t-1} on the right-hand side, i.e., the autoregressive specification; and, second, the possibility of cross-county error correlation. Let us consider each in turn. When the explanatory variables include the lagged dependent variable (Y_{t-1}) *and* the error term is characterized by autocorrelation, correlation between Y_{t-1} and the error term results. Without further action, parameter estimates will be inconsistent (Judge et al. 1985); typically, the effect of the lagged dependent variable will be exaggerated, because it captures the effects of omitted determinants. Two statistical solutions present themselves (Montgomery & Casterline 1993). Both solutions make recourse to the method of instrumental variables. In the first, additional lags in Z and X serve as instruments for Y_{t-1}. In a two-stage approach, the prediction equation for Y_{t-1} is first estimated (using the double transformation to correct for autocorrelation and heteroskedasticity described above), and then the equation for Y_t is estimated (again using the double transformation described above) with the instrument predicted from the first step substituted for the observed Y_{t-1}.

This solution yields consistent estimates if one is prepared to assume that the error term U_{it} is uncorrelated with Z and X. Suppose, however, that the error term is a composite:

$$U_{it} = v_i + w_{it} \tag{3}$$

where component w_{it} is unrelated to X and Z, and component v_i summarizes all time-invariant county-specific unobservables, which may or may not be related to X and Z. This specification of U_{it} essentially acknowledges that our set of X and Z variables is incomplete. If v_i is correlated with Z and X — and there is no reason to assume a priori that they are unrelated — then the effect of the lagged dependent variable Y_{t-1} continues to be exaggerated (i.e., effects of omitted determinants are attributed to Y_{t-1}) even under the instrumental variables estimation approach just described. (Using the estimated county-specific level of natural fertility as an explanatory variable is a step towards coping with the v_i, but it is an incomplete solution.) To remove v_i, all variables in equation 2 are reexpressed as deviations from county-specific means. This sweeps the fixed county effect v_i (representing all persistent county-specific variables) from the equation, as well as the time-invariant variables H and natural fertility. We then proceed with the two-stage instrumental variables approach. We term this estimation method the "instrumental variables/deviations" approach.

Turning to the estimation problems presented by cross-county error correlation, one can view this as reflecting correlation across counties in the fixed county effects v_i, i.e., $\mathrm{Cov}(v_i, v_j) \neq 0,\ i \neq j$. These can be described as persistent regional patterns of fertility, due to omitted determinants that themselves follow regional patterns. Such cross-county error correlation threatens the consistency of estimation of equation 2 because of the presence of $\ddot{Y}_{t-1}$ on the right-hand side (Anselin 1988; Cliff & Ord 1981; Doreian 1982). Our substantive concern again is that $\ddot{Y}_{t-1}$ will capture the effects of omitted determinants. But note that under the "deviation-instrumental variables" estimation approach all county effects v_i are eliminated, and thus regional correlations among the v_i do not intrude. Thus, our solution to this problem is also the deviation-instrumental variables estimation approach.

Results

Table 1 shows the means and standard deviations of the variables in the analysis. Note the structure of the data: 100 time series covering the period 1961-88, or 28 cross-sections for 100 counties. Some variables are time-constant, namely the three diffusion catalysts and the control variable "natural fertility." Family planning access (lagged 2 years) takes on nonnull values only in the period after 1969. Whereas some variables show dramatic changes over time (e.g., birth control and child mortality), other variables change rather slowly. To ease the interpretation of the interaction results, for the regression analysis the diffusion catalysts (H variables) are transformed into dummy (1, 0) variables, with their means serving as cut points.

Figure 3 summarizes the temporal and spatial pattern of the Costa Rican fertility transition. In Figure 3A, counties are shaded according to the period during which the fertility rate fell below a threshold of 0.350. In Figure 3B, the observed years are smoothed using the trend-surface technique (Cliff & Haggett 1988). (A cubic surface of the geographic coordinates of counties' central places is fit to the observed years, with an adjusted R^2 of 0.46.) The maps suggest that the transition began in the central and northeastern part of the country and then spread quickly through a corridor running parallel to the Pacific coast. In the trend-surface map, the first wave of fertility decline is just an extrapolation from the northeast Central Valley to sparsely populated areas. The first meaningful wave in the map (threshold years' 1965-66) encompasses the Central Valley and the northern half of Limón province in the east. During the two years following this wave, the fertility decline spread as far as 120 km to the west. These first waves cover vast areas of the country, an indication of rapid spread. After these early waves, however, the fertility transition moved much more slowly toward the north (toward Nicaragua) and east (Panama), advancing only about 10 km every two-year period. The narrow bands of progress in the 1970s suggest the existence of barriers to diffusion, which probably were of a cultural nature,

TABLE 1: Descriptive Statistics for Variables Used in the Analysis[a]

	1961-1988		1961-1964		1973-1976		1985-1988	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Dependent variable								
Y_t Birth control	475	234	146	139	588	145	645	109
Diffusion variables								
Y_{t-1} Lagged birth control	436	219	116	95	513	108	640	80
$\ddot{Y}_{t-1}$ Out-county birth control	460	209	99	61	571	55	642	36
Diffusion catalysts H								
Geographic compactness*	842	134	*	*	*	*	*	*
Social compactness*	555	108	*	*	*	*	*	*
Openness*	608	120	*	*	*	*	*	*
X_{t-2} Variables								
Socioeconomic development	482	130	379	111	492	116	561	117
Social security	367	253	140	143	329	191	693	115
Legal unions	839	155	860	163	838	153	818	150
Child mortality	124	81	205	63	123	50	36	20
Z_{t-2} Variable								
Family planning access	415	412	0	0	477	247	773	338
Natural fertility*	480	71	*	*	*	*	*	*
N	2,800		400		400		400	

[a] Pooled time series 1961-88, 100 counties. All figures per 1,000.
* Time-constant variable; 100 observations

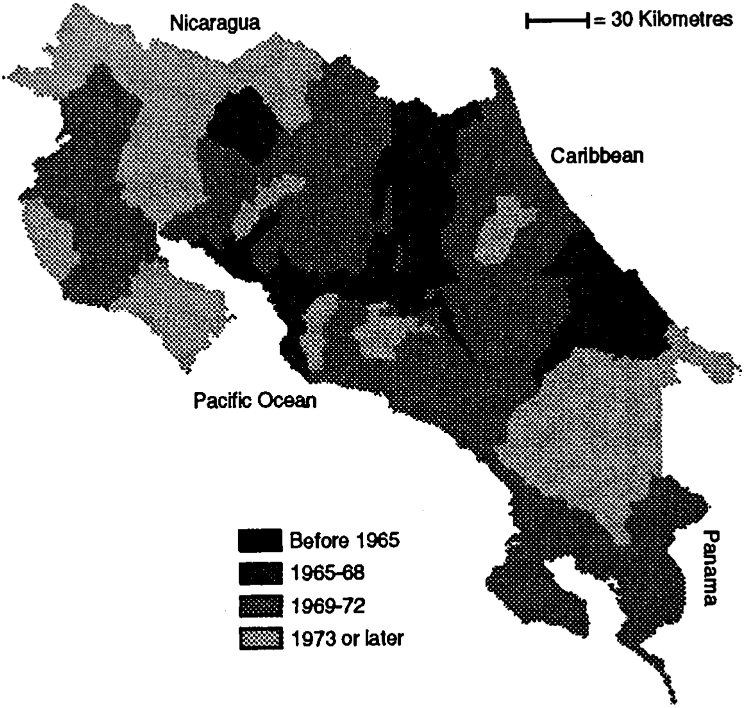
given that the population in those areas are in many aspects culturally closer to the neighboring countries than to the population in the Central Valley.

The spatial order evident in the maps suggests that interaction diffusion, through contact between adjacent areas, might have contributed to the timing and pace of the transition. The spatial pattern is not, however, conclusive evidence of diffusion. It may reflect nothing more than the spatial patterning of demand and supply factors. A more demanding test of the contribution of social contagion is provided by regression analysis that examines social interaction effects net of socioeconomic and family planning program variables.

We first consider estimates of the effects of the more conventional fertility determinants in an equation that does not contain diffusion effects. These effects can be estimated through the GLS procedure described above. The results from this preliminary analysis for the most part conform to expectations (Table 2). The estimated effects are conditional, however, on whether or not a linear time

FIGURE 3: Maps of the Year in Which Counties Surpassed the Threshold $F = 0.350$

A. Observed

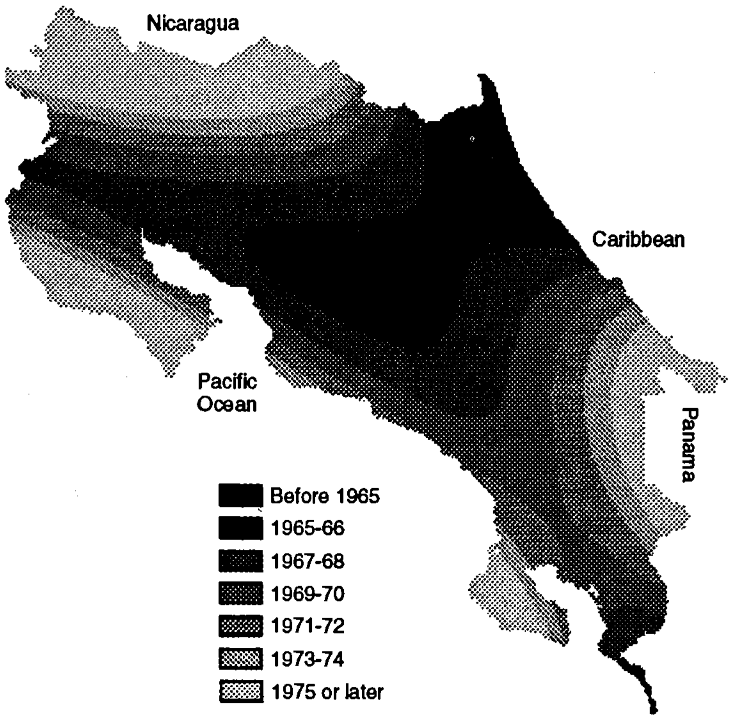


trend is included in the equation. Most of the effects are much smaller in magnitude (and in some instances reverse sign) when time is added to the equation; only the index of socioeconomic development is unaffected. This indicates that apparent effects of some socioeconomic variables are more properly attributed to other historical developments, such as the mass media, affecting the country as a whole. Without a control for time, all regression coefficients have the expected sign and are statistically significant (t ratio > 1.9).

Socioeconomic development shows the expected positive impact on birth control. The elasticity indicates that a 1% increase above the mean of socioeconomic development will produce a 0.38% rise in the birth control proportion. The proportion of the population covered by social security has contradictory effects on birth control: positive (as expected) without a control for time, negative with a control for time. The effect of child mortality is negative in both equations, but much weaker with a control for time: a 1% reduction from the mean child mortality rate results in a 0.06% increase in the birth control

FIGURE 3: Maps of the Year in Which Counties Surpassed the Threshold $F = 0.350$ (Continued)

B. Trend-surface smoothing



proportion. With a control for time, the effect of legal unions is not significant (and in direction counter to expectations).

The demographic effect of the family planning program is difficult to assess empirically. The program started in 1968 when the fertility transition was already well under way. In addition, the program quickly covered most of the nation, undermining the validity of areal analyses that must maintain a distinction between program and nonprogram areas. Therefore, it is not surprising that the positive and significant effect of family planning access in Table 2 disappears with a control for time. Moreover, the significant positive effect without the control for time is modest: a 1% increase above the mean accessibility would result in merely a 0.07% increase in the birth control proportion.

As expected, the proxy for natural fertility is inversely related to birth control. As indicated above, this relationship is probably not causal but rather is an indication of bias in the measure of birth control adoption: our measure

TABLE 2: A Nondiffusion, Multiple Regression Model on Birth Control Proportion[a]

	No Time-trend Control			Time-trend Control		
	Coef.	t Ratio	Elasticity at Means	Coef.	t Ratio	Elasticity at Means
X_{t-2} Variables						
Socioeconomic development	0.39	6.8	0.39	0.38	7.1	0.38
Child mortality	-1.17	-15.4	-0.30	-0.23	-2.5	-0.06
Social security	0.24	9.1	0.19	-0.11	-3.4	-0.08
Legal unions	-0.21	-7.2	-0.37	0.05	1.6	0.09
Z_{t-2} Variable						
Family planning access	0.08	5.9	0.07	0.01	1.0	0.01
Natural fertility	-0.29	-4.4	-0.29	-0.50	-7.1	-0.51
Time trend (60 - 1960)	—	—	—	0.02	20.7	2.97
Constant	0.61	12.6	1.30	-0.87	-10.1	-1.84
Buse R^2	0.63			0.65		
Average ρ	0.80			0.79		

[a] Pooled time series 1961-88, 100 counties
Estimation by generalized least squares (heteroskedastic, autoregressive model). See text.

underestimates the true level of adoption in counties with lower natural fertility. Inclusion of the natural fertility proxy is intended to control for this bias.

We now turn to the estimated diffusion effects. Table 3 presents the full equation estimated through the "instrumental variables/deviations" approach described above. The diffusion effects — both within-county and intercounty — are positive in sign and statistically significant (at the 0.001 level). This constitutes the principal empirical result of this research, and thus we should pause briefly, before getting further tangled in our interpretation of this result, to underscore its importance. (Results from the more naive estimation approaches termed GLS and instrumental variables are not presented, but it is important to underscore that their diffusion effects were substantially stronger and consistent than those from the statistically soundest approach presented in Table 3.) We have hypothesized that social interaction processes affected the timing and pace of the Costa Rican fertility transition. Lacking direct microlevel measures of social influence effects, we search for evidence suggestive of these effects in the spatial-temporal pattern of fertility change. Table 3 contains just such evidence: fertility in a county is strongly associated with prior levels of fertility in the same and adjacent counties, net of measured variables and unmeasured fixed county effects.

Recall that to evaluate the within-county diffusion effects both Y_{t-1} and squared Y_{t-1} must be taken into account. The positive sign on the first term and the negative sign on the second indicates that within-county diffusion effects are stronger early in the fertility transition, i.e., when the prevalence of birth control is low. The declining importance of diffusion in part reflects the exhaustion of the pool of potential adopters, but it may also reflect the greater salience of the social interaction mechanisms described above — information flow, demonstration effect, change in normative context — in the early stages of transition. As the transition progresses, the information flow and demonstration effect may shift to external agencies (e.g., family planning program information, education, and communication [IEC] efforts, medical counseling).

While the diffusion effects shown in Table 3 provide overall confirmation of the hypotheses motivating this research, the point estimates themselves are not entirely interpretable. The point estimates of the within-county diffusion effects indicates that these effects disappear when the proportion practicing birth control reaches about 0.40, which seems a bit low. Moreover, the estimates suggest that when the proportion practicing birth control reaches a high level, a negative feedback sets in; i.e., increased prevalence of birth control in itself stimulates a countervailing reduction in prevalence. While negative feedback is not implausible — awareness of unhappy experiences may spread through social interaction — we do not think that much weight should be placed on this result. The quadratic is a relatively crude functional form and inherently unstable for point estimates in the tails of a distribution. The procedure for estimating the proportion practicing birth control is also crude (a linear transformation of the marital fertility rate). Finally, specification error might also contribute to this result, since no effort was made to include in the equation variables that explain the fertility boom that took place circa 1980 in the more modern strata of the Costa Rican population. This may make it appear as if high levels of birth control had a negative, autoregulating feedback.

Comparison of the inter- and the within-county effects also indicates that too much weight should not be placed on the point estimates. This comparison indicates that the intercounty effect exceeds the within-county effect whatever the prevalence of birth control. We expect that, in general, within-county diffusion effects should exceed intercounty effects, under the assumption that social interaction is more intense within counties. By this reasoning the relative magnitudes of the two effects are reversed, suggesting that either the point estimates of the within-county effects are too small or the point estimates of the intercounty estimates are too large. We lean towards the latter view. Of course it is not implausible — although it seems unlikely — that a mechanism such as information flow is more active across county boundaries than within counties.

The regression in Table 3 offers separate estimates of within-county diffusion effects for counties that are low and high on social and geographic compactness. Neither of the pertinent interactions tests are significant, however.

TABLE 3 "Instrumental Variables/Deviations" Estimate of the Regression Coefficients of the Model on Birth Control Proportion, with Diffusion Effects[a]

	Coef.	t Ratio
Within-county diffusion		
Y_{t-1} Lagged birth control	0.35	6.7
Y^2_{t-1}	-0.40	-4.8
$H\,Y_{t-1}$ Interactions with:		
Social compactness	0.00	0.0
Geographic compactness	-0.03	-0.9
Intercounty diffusion		
$\ddot{Y}_{t-1}$ Out-county birth control	0.50	11.8
$H\,\ddot{Y}_{t-1}$ Interaction with openness	-0.04	-1.8
X_{t-2} *Demand variables*		
Socioeconomic development	0.26	3.9
Child mortality	-0.13	-1.8
Social security	0.03	1.0
Legal unions	-0.09	-1.2
Z_{t-2} *Supply variable*		
Family planning access	-0.02	-2.2
Time trend (60 = 1960)	-0.001	-0.8
Constant	0.07	-1.0
Buse R^2	0.79	
Average ρ	0.55	

[a] Pooled Time Series 1961-1988, 100 counties

Estimate approach using instrumental variables for lagged birth control and data transformed into deviations from county's means.

Similarly, the intercounty diffusion effects differ very little between counties rated less and more open. These Costa Rican data do not support the hypotheses that diffusion effects are conditional, to a meaningful extent, on the internal homogeneity of an area or on the openness of the area to other areas. It may be that our measures of homogeneity and openness are simply too crude for testing these hypotheses.

Regarding the conventional demand and supply variables, it is useful to compare the estimates in Table 2 with the net-of-diffusion effects in Table 3. Adjustment for diffusion effects reduces substantially the net effects of socioeconomic development, child mortality, and social security. All three show significant net effects without controls for diffusion effects, whereas only socioeconomic development shows significant net effects once diffusion effects are controlled. The prevalence of legal unions is not significant whether or not diffusion effects are controlled (if time is controlled). Access to public family planning facilities shows a perverse negative effect once diffusion effects are controlled. This effect points out again the difficulties of evaluating the impact of the Costa Rican family planning program. These results, although based on a small set of conventional variables, suggest that diffusion dynamics account for a large portion of the apparent effects of conventional variables on the Costa Rican fertility transition. There are two alternate interpretations of this outcome: (1) Social contagion was a key *mechanism* through which conventional variables acted upon fertility, or (2) most of the apparent effect of conventional variables was, in reality, a reflection of social interaction diffusion dynamics.

Discussion

Most of the existing theory and empirical research on fertility transition implicitly assumes isolated individuals: people weighing privately the benefits and costs of children and the objective and subjective costs of birth control to make reproductive decisions. The adherence to routine behaviors dictated by cultural norms is also sometimes acknowledged. From this standpoint, the fertility of a population is a simple aggregation of individual motivations and cost constraints. This is a model of fertility behavior that is in contradiction to basic sociological theory and evidence about interdependencies among individuals and the existence of powerful social influence effects.

In this article we relax the isolation assumption, allowing reproductive behavior to be influenced by social contagion. The central hypothesis, stated simply, is that adoption of birth control by some individuals affects the probability of adoption by others. The implication is that the reproductive behavior of a poorly-educated farm couple, for example, will be different if those with whom they interact are practicing birth control than if they are not.

As noted at the outset, attention to social influence effects is hardly novel in sociology. Even in research on fertility transition, that the practice of birth control "diffuses" from an elite to the masses has been noted, as a descriptive given, from the earliest literature on fertility decline to the present. And in the last two decades, studies of contemporary and historical fertility transitions have, with increasing frequency, invoked the notion of diffusion to explain puzzling features of the spatial-temporal patterns of fertility change. Our objective in this article is to add rigor to the ongoing debate by focusing on social interaction diffusion and by conducting empirical tests for diffusion effects on the timing and pace of the Costa Rican fertility transition.

The Costa Rican transition possesses many of those features thought to be suggestive of diffusion dynamics: the decline occurred extremely rapidly, outpacing socioeconomic development and reaching all major strata of the population in a relatively brief period of time; there is no evidence of a downward shift in family size preferences that corresponds, either in timing or in magnitude, with the fertility decline; and the spatial-temporal pattern of fertility decline shows a spatial ordering suggestive of contagion between neighboring areas. To this evidence we have added regression estimates in which social interaction diffusion effects — endogenous feedback effects — are quantified net of the effects of other socioeconomic, cultural, and family planning program factors, both observed and unobserved. Our aim is to put the diffusion thesis to a focused and demanding empirical test.

The regression analysis supports the hypothesis that interaction diffusion dynamics affected the course of fertility transition in Costa Rica. According to the regression estimates in Table 3, diffusion dynamics substantially accelerated fertility decline in Costa Rica, especially in its early stages. The evidence concerning the catalyzing role of the compactness and openness of the county, in contrast, is less persuasive.

We conclude with brief comments about the validity of the results, their theoretical significance, and their policy relevance.

With respect to the validity of the results, we have already discussed at length the risk that estimated social interaction effects in fact reflect the action of variables omitted from the analysis. There can be little doubt that, above and beyond measured socioeconomic and programmatic conditions, fertility change in Costa Rican counties was associated with the number of people controlling their fertility in the county and in surrounding counties. To minimize the possibility that this is the consequence of omitting important explanatory variables, we have taken care in developing the statistical model. Nevertheless, it must be recognized that, in the absence of experimental design (far-fetched in this instance), we must live with this possibility. We are also not entirely comfortable with analysis at the aggregate level. The social dynamic hypothesized is intrinsically multilevel: specific individuals are influenced by the behavior of others, possibly a single other person but more plausibly the aggregation of observed behaviors and/or attitudes of others; the resulting changes in the behavior of specific individuals, in turn, modifies the aggregate behavior that other individuals encounter (Coleman 1990). Multilevel analyses are clearly in order. More to the point is social network research that measures and models social interaction processes in relation to reproductive behavior (Casterline & Knight 1993; Watkins 1990).

An important theoretical question is whether birth control contagion — assuming that it does exist — is a distinct causal dimension or just a mechanism for the action of more conventional fertility determinants. Social interaction diffusion effects, as we have defined them, cannot be captured through recruitment of additional explanatory variables, i.e. as further entries in standard diagrams of static causal relationships with their own boxes and corresponding arrows. Interaction diffusion essentially is a formal feedback that only becomes apparent in dynamic graphs. Diffusion effects do not operate at equilibrium, the state that static models usually assume (Rosero-Bixby & Casterline 1993). In this respect diffusion effects of the sort we have specified here are a subclass of the broader class of social influence effects that have been of interest to sociologists for some time. A peculiarity of the diffusion effects is that they cannot act alone: an innovation, a pool of susceptibles, and a rate of social interaction are needed for diffusion effects to occur. They are wrapped up in the *process* of fertility change; they cannot be extricated, theoretically or empirically, from the process itself. If diffusion effects as we have defined them cannot be identified apart from fertility change, is it reasonable to view them as making a causal contribution to fertility change?

We answer this question in the affirmative, provided that one is assessing causal contribution to the timing and pace of transition. One can construct precise hypotheses about causal impacts of diffusion dynamics on the timing and pace of transition and, as illustrated in the empirical research presented in this article, these hypotheses are amenable to testing with existing data. Nevertheless, it is clear that diffusion effects are different in nature than those of traditional fertility determinants, and we confess to some unease with the existing theoretical treatment of diffusion effects, in our work and in the work of others.

The payoff from viewing diffusion as a distinct causal dimension becomes clearer when one turns to policy considerations. The planning, and evaluation, of family planning programs can be seriously biased if one fails to consider the potential multiplier effects of diffusion dynamics. Indeed, diffusion dynamics can be exploited on behalf of population policy, for example by designing programs that reduce the taboo character of reproduction topics or that facilitate social interaction about reproductive behavior (contraception in particular). The family planning program in Taiwan in the 1960s, for example, quite consciously took advantage of diffusion dynamics (Freedman & Takeshita 1969).

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