The Population Ecology of Organizational Founding: Location Dependence and Unobserved Heterogeneity

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Using pooled cross-sectional time series data collected at the level of 13 geographical regions, this paper investigates the effects of location dependence and unobserved heterogeneity on founding rates of rural cooperative banks in Italy from 1964 to 1988. Parametric and semiparametric models of organizational founding reveal that the organizational population is internally differentiated and that different segments of the population respond heterogeneously to general processes of legitimation and competition. This study emphasizes the importance of identifying the correct level of analysis at which population-level processes operate and of accounting for unobservable factors related to the cross-sectional structure of organizational populations. The findings demonstrate how an ecological approach that incorporates information on the spatial structure of the population can provide a more detailed understanding of the evolutionary dynamics of organizations.•

Organizational environments have spatial components that affect the evolutionary dynamics of organizational populations. First, geographical barriers of various kinds may allow enough isolation for different evolutionary paths to be explored in different regions (Eldredge and Gould, 1972; Mayr, 1976). Second, localized resource environments may pose complicated problems of adaptation for individual organizations (Carroll, 1985; Hannan and Freeman, 1989; Baum and Mezias, 1992). Finally, processes of legitimation and competition, which respond to organizational density, may vary depending on the geographical boundaries used to define organizational populations (Carroll and Wade, 1991; Hannan and Carroll, 1992). When taken together, the consequences of heterogeneity in the spatial distribution of resources have far-reaching implications for the ecological dynamics of organizational populations, because the level of spatial aggregation defines implicitly the population boundaries (Singh, 1993), and the intensity of competition among organizations is a function of the similarity in resource requirements (Hannan and Freeman, 1989; Baum and Mezias, 1992).

Based on the argument that markets and other institutional arenas eventually evolve at least to the national level, organizational ecologists usually specify population processes at that level (Hannan and Carroll, 1992). The propagation of organizational populations in space to the national and often supranational level, however, does not imply spatial homogeneity. Many organizations remain local and depend heavily on their immediate institutional and competitive environments for support, resources, and demand. But location dependence is not a phenomenon restricted to small or marginal organizations. The most striking feature of economic and organizational activity is their geographical concentration: Certain areas become so highly specialized that production in many industries is concentrated in space (Krugman, 1991a). It has frequently been mentioned, for example, that "Nighttime satellite photos of Europe reveal little of political boundaries but clearly suggest a center-periphery pattern whose hub is

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somewhere in Belgium" (Krugman, 1991b: 484). Similarly, the manufacturing *belt* in the U.S., the tile *cluster* and the textile *district* in northern Italy, the research *triangle* in North Carolina, the machine-tool *district* in Germany, the *networks* of suppliers surrounding the *kaisha* (lead manufacturers) in Japan, Silicon *Valley,* and Route 128 are all expressions that conjure up images of highly localized organizational activities (Piore and Sabel, 1984; Best, 1990; Porter, 1990; Preer, 1992).

The recurrence of patterns of organizational concentration in space across different industries and in a number of national contexts provides indirect evidence that location may be a general factor shaping the evolution of organizational populations. If forces exist that give evolutionary advantages to organizations located near other organizations or in specific geographical areas, then the internal structure of the organizational population can no longer be considered homogeneous, and organizational birth and death rates will vary systematically across locations. Also, if all organizations in a population do not compete for the same scarce resources or contribute to competition in the same way, then it is essential to select the appropriate level of analysis to examine the dynamics of the population, since different levels of spatial aggregation imply different assumptions about how general processes of legitimation and competition unfold.

This paper pursues these issues of location dependence and unobserved heterogeneity in the context of processes of organizational creation, for two reasons. First, in spite of the accumulation of empirical results in the population ecology of organizational founding, relatively little is known about the consequences of the level of spatial aggregation on founding rates in organizational populations. Lacking explicit theoretical indications, various researchers have tried different levels with different results (Barnett and Carroll, 1987; Carroll and Wade, 1991; Swaminathan and Wiedenmayer, 1991; Baum and Mezias, 1992; Baum and Singh, 1994; Hannan and Carroll, 1992; Freeman and Lomi, 1994). In their analysis of the U.S. brewing industry, for example, Carroll and Wade (1991) found support for ecological models of density dependence at the state and regional levels but not at the city level, while Carroll and Hannan (1989) found that founding rates of newspaper organizations depend on density in a way that is consistent with ecological theories only in small metropolitan areas. Second, while problems related to unobserved heterogeneity have been raised frequently in the context of processes of organizational mortality (Freeman, Carroll, and Hannan, 1983; Hannan, 1988; Petersen and Koput, 1991), virtually no research has been done on the implications of unobserved heterogeneity for organizational founding rates.

These two problems are distinct but are conceptually related. The level of aggregation problem has its origins in the fact that ecological theories are silent about how the boundary of a population should be drawn and thus at which level general population processes of legitimation and competition actually operate to shape the vital dynamics of organizational populations (Singh, 1993). The problem of heterogeneity in the study of organizational founding rates originates from the logical impossibility of associating organization-level attributes to organizations whose appearance is to be recorded and from the methodological difficulty of dealing with non-events (Delacroix and Carroll, 1983; Hannan, 1989, 1991; Barron and Hannan, 1991). In studies of organizational founding, it is often not clear what selection forces are operating on (Singh, 1993), and sources of unobserved heterogeneity have to be sought at the population level (Singh and Lumsden, 1990). But this brings us back to the first problem: Which population level?

To address these conceptual issues in the population ecology of organizations, I propose a specific model of location dependence in organizational founding rates. I then test the implications of the model using data on the founding of Italian rural cooperative banks during the period 1964–1988. Italian rural cooperative banks are especially useful for this analysis because (1) the institutional and competitive environments of this population are known to be highly heterogeneous by region, (2) location is of crucial importance to these specialist banks because they cannot expand by branching beyond the local level, and (3) accurate and complete data exist, allowing the effects of location dependence and heterogeneity on founding rates to be tested. The population selected for this study provides an opportunity to evaluate the effects on organizational founding rates of unobservable components related to spatial heterogeneity in the distribution of resources and to learn more about the level at which competitive and institutional forces unfold to shape the evolution of organizational populations. Because of the role that these organizations play in the economy and society, studying the population ecology of credit institutions also has considerable interest in its own right. This study also adds to the findings of others who have studied their population dynamics (Ranger-Moore, Banaszak-Holl, and Hannan, 1991; Barron, 1992a; Rao and Neilsen, 1992; Haveman, 1993; Amburgey, Dacin, and Kelly, 1994; Freeman and Lomi, 1994). This study of banks should also clarify the organizational dynamics of many other societal sectors that similarly have both technical and institutional characteristics (Scott and Meyer, 1983; Powell, 1991).

THEORETICAL BACKGROUND

Processes of natural selection affect the distribution of organizational populations in time and space. Evidence for selection can therefore be inferred either from changes over time or from spatial variation (Manly, 1985). While organizational ecologists have emphasized the effects of time heterogeneity and time dependence in the evolution of organizational populations (Hannan, 1988; Hannan and Freeman, 1989: chap. 5), issues of spatial heterogeneity and location dependence have remained unexplored until very recently (Carroll and Wade, 1991; Hannan and Carroll, 1992). This is surprising for a research program that has its intellectual roots in human ecology, in which spatial processes related to community succession have played a prominent role (Park, 1936; Hawley, 1950). This neglect is even more surprising if one considers that whenever assumptions of population homogeneity with respect to location are relaxed, a number of conceptual issues surface to complicate the ecological analysis of organizational founding (Hannan and Carroll, 1992: chap. 7).

Location Dependence in Organizational Founding Rates

Much recent ecological research focuses on the relationship between density (the number of organizations in a population) and organizational birth and death rates (Hannan, 1986; Hannan and Carroll, 1992). The theory of density-dependent organizational evolution postulates that opposing processes of legitimation and competition shape organizational birth and death processes and that legitimation and competition are systematically linked to density. According to the model implied by this theory, founding rates in organizational populations are directly proportional to the legitimacy of the organizational form and inversely proportional to the intensity of competition. At low levels of density, legitimation processes dominate, causing founding rates to rise and mortality rates to decline, while the opposite happens at high levels of density. Empirical research has shown that the founding rate has a nonmonotonic, inverted U-shaped relationship with density, a result that is usually interpreted as supporting the theory (Hannan, 1991; Hannan and Carroll, 1992).

In the most comprehensive study to date, Hannan and Carroll (1992: chap. 4) reported results that support the theory of density-dependent founding rates in four of the seven populations studied, but for organizational mortality, the effects of density were always found to be consistent with the theoretical predictions (Hannan and Carroll, 1992: chap. 6).[1] Because one of the strengths of density dependence theory is its symmetry, i.e., that it applies equally well to processes of organizational founding and disbanding (Hannan, Barron, and Carroll, 1991), it is particularly important to understand sources of potential anomalies.

One explanation for findings inconsistent with the theory of density-dependent founding rates is based on the choice of the level of analysis (Carroll and Hannan, 1989; Carroll and Wade, 1991; Hannan and Carroll, 1992: chap. 7). If founding rates in organizational populations are location-dependent, the cross-sectional structure of the population is likely to interact with longitudinal processes of change and produce heterogeneous local responses to general processes of legitimation and competition. Zucker (1989: 543) conjectured that "smaller geographical areas should theoretically involve more intense competition since they are tightly bounded resource arenas." Similarly, Carroll and Huo (1986: 845) studied the evolution of newspapers in the San Francisco Bay area because "American newspaper organizations depend on markets bounded by metropolitan areas . . . [which represent] . . . relatively autonomous organizational environments." More recently, Carroll and Wade (1991: 272) hypothesized that "legitimation and competition manifest themselves more strongly at different levels of analysis for mortality and founding. For instance, founding processes may be more localized."

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In one case, however, the second-order effect turned out not to be statistically significant.

A related issue is that at lower levels of spatial aggregation, such as the regional level, local (sub)populations may have later starting dates than the population as defined at a higher level, like the national level. The observable cross-sectional configuration assumed by a population at a given time reflects, in part, early patterns of proliferation and diffusion of the organizational form (Stinchcombe, 1965). One cannot assume, however, that later subpopulations start from scratch in terms of their legitimation, especially if the form is already taken for granted (Singh, 1993). Carroll and Wade (1991) encountered this problem in their study of the American brewing industry, since there is heterogeneity among the states as to when prohibition actually began to be enforced. Differences in legislation across states are a common source of location dependence. Thus an analysis of organizational vital rates conducted at the regional level is likely to miss the initial legitimation effects of density except in the region where the population first started, while an analysis conducted at the national level will be based on data that have been aggregated over cross-sectional units that are heterogeneous in terms of basic institutional and competitive processes.

The key point to this discussion is that shifting the level of aggregation involves redefining the population boundaries, because the level of analysis implicitly determines the relevant context for legitimation and competition processes (Singh, 1993). Unfortunately, population ecology theories of organizations offer no specific guidance on how to choose among alternative, but equally plausible population boundaries (Carroll and Wade, 1991). Choosing the appropriate level is crucial to test ecological theories of organizational founding because density acts in different ways across levels of analysis. It is plausible to think that potential founders will be more sensitive to local variations in the levels of legitimation and competition because of limits in their capacity to collect information on nonlocal resource conditions and because of the ambiguity involved in interpreting events taking place in more distant sites. For these reasons, density dependence in organizational founding rates is expected to operate more strongly at lower levels of analysis and the model of density-dependent founding rates to be better specified at the local than the national level of analysis.

Unobserved Heterogeneity in Organizational Founding Rates

Investigating the role of location in the relationship between density and organizational founding rates directs attention to issues of heterogeneity within organizational populations (Haveman and Romanelli, 1994). Heterogeneity is a conspicuous problem in the population ecology research program because homogeneity is usually assumed in analyzing organizational populations as meaningful social entities (Hannan and Freeman, 1989) and because unobservable factors related to location may affect the estimates of theoretically important structural parameters (Petersen and Koput, 1991; DiMaggio, 1994). This problem is particularly vexing in the study of organizational founding because organizational attributes cannot be used as independent variables, since there is no organization prior to founding (Delacroix and Carroll, 1983: 275). If founding rates are location-dependent, then aggregation across spatially heterogeneous units will introduce unobservable components into the sample, which, if not properly accounted for, will result in specification bias (Hsiao, 1986).

Issues of unobserved heterogeneity have typically been addressed in studies of organizational mortality, in which organization-level data are available and it is easier to test how selection processes operate on individual units as a function of their age, size, evolutionary strategy, and linkages with other organizations (Freeman, Carroll, and Hannan, 1983; Carroll, 1985; Hannan, 1988; Barnett and Amburgey, 1990; Baum and Oliver, 1991; Swaminathan and Wiedenmayer, 1991; Haveman, 1992). In this context, unobserved heterogeneity is usually related to the omission of organization-specific variables, which produces spurious density dependence (Petersen and Koput, 1991).

Studying unobserved heterogeneity in processes of organizational founding is more complicated because it is less clear exactly what selection forces are operating on, since organization-level characteristics cannot be observed for units not yet arrived in the population. An example from current research on localized competition may help to illustrate this point more vividly. In their study of the Manhattan hotel industry, Baum and Mezias (1992) found that hotel failure rates were increased significantly by the effects of localized competition in terms of geographic location, size, and price. Localized competition was measured by comparing the position of a focal hotel to the position of others within a given distance. They found that more similarly sized, priced, and located hotels compete more intensely. No effect of location on organizational founding rates could be studied in this framework because distances cannot be computed between attributes of organizations that do not yet exist. The conclusion is that in studying processes of organizational founding, the population itself must be seen as the unit experiencing the events (Hannan, 1989), and sources of unobserved heterogeneity have to be sought at the population level. But this brings the argument back to location dependence, since populations can be studied at different levels of analysis.

The problem of studying founding as occurring in a homogeneous population context is that not all potential founders are equally at risk of starting organizations or equally able to take advantage of local opportunities to mobilize resources, because exposure to information and availability of opportunities vary significantly across space and time. Variations in social and economic conditions across sites or local differences in the population's institutional history will produce differences in intrinsic founding rates, or in unobservable region-specific "proneness" to experiencing the founding of particular organizational forms. But if different segments of the organizational environment cannot be considered equally at risk of experiencing the founding of an organization of a given type, organizational populations can be expected to be partitioned into discrete segments, or latent classes, that respond heterogeneously to general processes of legitimation and competition. For these reasons, the dependence of organizational founding rates on density is expected to vary significantly across these heterogeneous segments.

The study of Italian cooperative banks allows us to address these concerns empirically in the context of an organizational population that offers an almost unique opportunity to explore the ecological consequences of location dependence and heterogeneity for organizational founding rates. Italian regions are characterized by a clear structural duality in economic and social conditions. For example, if we let the average gross domestic product (GDP) per capita in the European Union be equal to 100, the richest country is Luxembourg (125), the poorest Portugal (53.7), while Italy ranks seventh (100.4) after the Netherlands and before Belgium. Regional-level analysis reveals, however, that the average GDP per capita in the southern Italian regions is 70.7 (with a minimum in Calabria of 58.7), while the average GDP per capita in the northern Italian regions is 124, with Lombardia, Val d' Aosta, and Emilia-Romagna well above Luxembourg. This structural duality reflects and is a reflection of profound interregional differences in terms of employment (7 percent unemployment rate in the center-north against 23 percent in the south), composition of the labor force (8 percent of the labor force in the center-north is employed in the agricultural sector as opposed to 17 percent in the south), credit conditions (short-term interest rates in the center-north are about 3 percent lower than in the south), and investor preferences (more than 50 percent of the total wealth of northern Italian families is invested in stocks and bonds, while southern Italian families invest about 25 percent of their wealth in postal bonds and only 20 percent in stocks) (Galli and Onado, 1990).

Location is important because structures, processes, and functions of organizational populations are defined in and by time and space (Hawley, 1986; Dendrinos and Sonis, 1990), and evidence from previous studies suggests that organizational vital rates vary systematically across locations in a number of diverse organizational populations (Barnett and Carroll, 1987; Carroll and Wade, 1991; Swaminathan and Wiedenmayer, 1991; Baum and Mezias, 1992; Hannan and Carroll, 1992; Carroll et al., 1993). For these reasons, the present study of Italian mutual banks can contribute new evidence to this important stream of organizational research and illuminate some of the core theoretical questions in the ecology of organizational founding.

METHODS

Research Design and Data Structure

The population of Italian rural cooperative banks is especially suitable for the purposes of this study for three reasons. First, the well-documented heterogeneity in social and economic conditions across different geographical regions provides a unique opportunity to evaluate the effects of unobservable components related to location dependence on organizational founding rates (Brusco, 1982; Bagnasco, 1984; Piore and Sabel, 1984; Trigilia, 1986; Lazerson, 1988; Best, 1990; Perrow, 1992). Second, location is of crucial importance to these specialist banking organizations because they cannot expand by branching beyond the local level and are therefore very sensitive to localized processes of legitimation and competition (Costi, 1986; Lomi, 1995). Finally, the availability of accurate and complete data at the regional level permits the analysis of intertemporal variation in founding rates while controlling for observed and unobserved differences among cross-sectional units.

The study is based on pooled cross-sectional time series data collected at the level of 13 geographical regions and containing information on the founding of rural cooperative banks in Italy from 1964 to 1988. Of the 13 regions in the sample, four are located in the north (Friuli, Trentino, Veneto, and Lombardia), three in the center (Emilia-Romagna, Marche, and Lazio), five in the south (Abruzzi-Molise, Campania, Puglia, Basilicata, and Calabria), and one is an island region (Sicilia). The sample includes 91 percent of all the rural cooperative banks (RCBs) in the country. This percentage remained stable during the observation period (90.8 percent in 1964 and 91.4 percent in 1988). Italy is divided into 20 geographical regions, and there is no reason to believe that the thirteen regions selected are not representative of the overall situation. In 1964, RCBs represented 59 percent of the total number of banks in the country and 62 percent of the total number of banks in the sample. At the end of the observation period, in 1988, RCBs represented 66 percent of the total number of banks in the country and 68 percent of the total number of banks in the sample.

Information on founding activity was coded from various years of the official publications of the Banca d' Italia, *Bollettino Statistico della Banca d' Italia* and *Bollettino di Vigilanza,* and from the volume *Struttura funzionale e territoriale del sistema bancario italiano,* published by the Banca d' Italia (1977). Information on covariates was found in the *Annuario Statistico Italiano* (various years), published by the Italian National Institute of Statistics (ISTAT), and in the *Statistiche della Cooperazione* (various years), a statistical bulletin published by the Italian Ministry of Labor and Social Security.

Italian rural cooperative banks. Italian RCBs are unit banks specializing in financial services to local agricultural and craft businesses. The members of cooperative or mutual banks are also the main beneficiaries of the financial services rendered by the organization (Hansmann, 1988; Amburgey, Dacin, and Kelly, 1994). This distinguishes mutual from stock banks, which are owned by the stockholders and are presumably run to benefit those stockholders (O'Hara, 1981). As in the U.S. (Ranger-Moore, Banaszak-Holl, and Hannan, 1991), Italian legislation has maintained a sharp distinction between mutual and stock banks in the interest of protecting different types of customers. Mutual banks are owned by depositors whose main goal is to gain access to the financial services that the organization can offer. Stock banks are owned by stockholders whose goal is to maximize the return on their capital investment, and depositors are merely customers (Rasmusen, 1988). The charter of mutual and cooperative banks protects the organizational members right to have equal access to the product sold or services rendered by the organization. The charter of stock banks protects third parties and stakeholders against managerial abuses and ineptitude (Hansmann, 1988).

These basic differences between stock and mutual banks in terms of ownership, capital structure, and mechanisms of corporate governance are reflected in a number of normative constraints that severely limit the range of strategies available to individual organizations. For example, the law establishes that dividends paid to members cannot exceed the legal interest rate and must be proportional to the share of capital actually contributed, at least four-fifths of the total members must be farmers or craftsmen living or having their business in the same geographical area in which the bank operates, and financial operations with nonmembers cannot exceed 25 percent of the total deposits. Finally, mutual banks cannot expand by branching and cannot be changed into or be absorbed by private stock banks.

Rural cooperative banks have been the only partial exception to the trend toward a general reduction in the number of banks in Italy, as in many other industrialized economies, to alleviate problems of overbanking inherited from the financial crises of the 1920s and 1930s (Carenza, Frasca, and Toniolo, 1986). During the 1950s and '60s, policy makers and central credit authorities chose cooperative banks to diffuse banking and financial services in a country characterized both by strong regional imbalance in terms of the diffusion of industrial activities (Bagnasco, 1984) and by a complex texture of small and craft-like business firms that is still typical of the Italian economy (Goodman, 1989). Founding activity of cooperative banks peaked in the 1960s and remained significant through the '70s and '80s.

The diffusion of cooperative banks at the fringes of market areas occupied by the core national banks was a direct consequence of the policy efforts to control concentration in the industry and diffuse banking in market areas in which private stock banks could not operate profitably and were therefore unwilling to establish new branches. Strong differences in economic and social conditions across Italian regions have been one of the main reasons leading to the adoption of the cooperative organizational form to diffuse banking and financial services into peripheral and rural areas. Credit authorities used the creation of banks where organizational members are also the main beneficiaries of the financial services rendered by the organization as an organizational and policy instrument to reduce the imbalance in the geographical distribution and accessibility of financial services in the country without triggering direct competition at the industry level. Rasmusen (1988) proposed that mutual banking organizations in the U.S. may respond positively to conditions of relative economic backwardness and uncertainty, a hypothesis that is also suggested by observations of other kinds of cooperative organizations in the U.S. and abroad (Staber, 1989). Barnett and Carroll (1987), for example, found that mutual telephone companies were localized and proliferated in rural and relatively peripheral areas, while commercial companies occupied urban centers. Similar conclusions have been reached by Ben-Ner (1987) in an article surveying the literature on producer cooperatives.

Figure 1. Rural cooperative bank foundings by year.

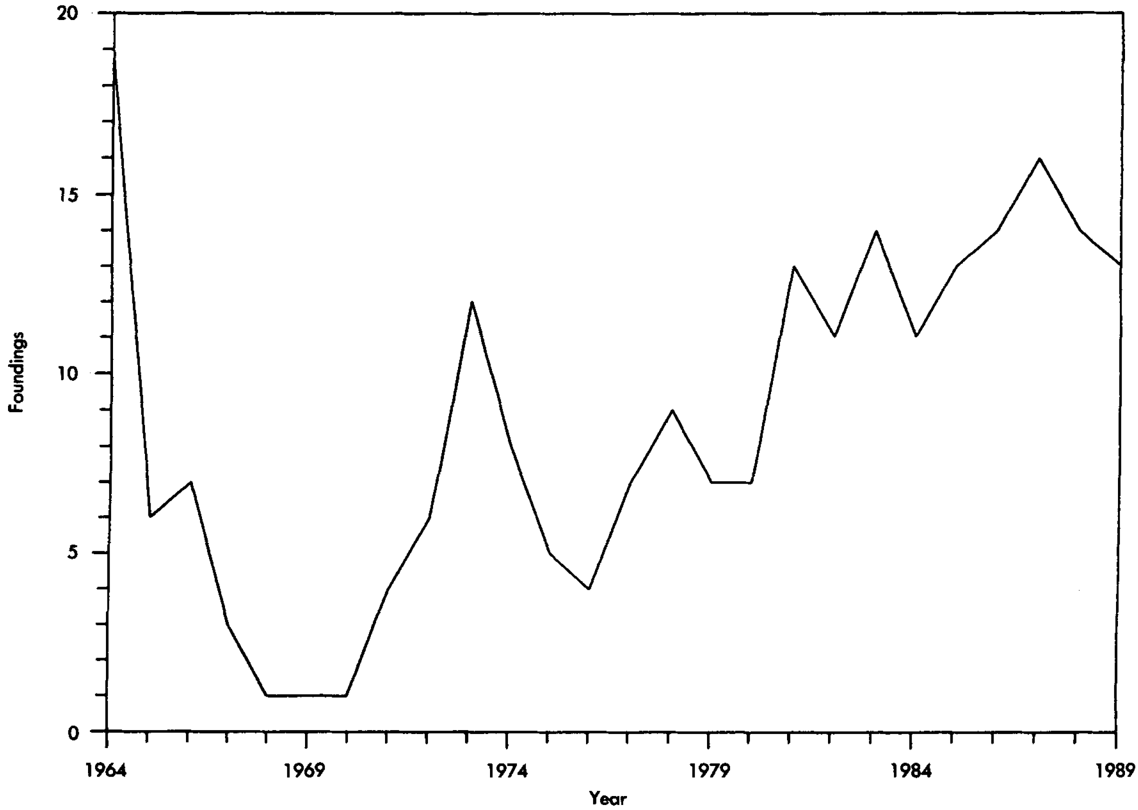
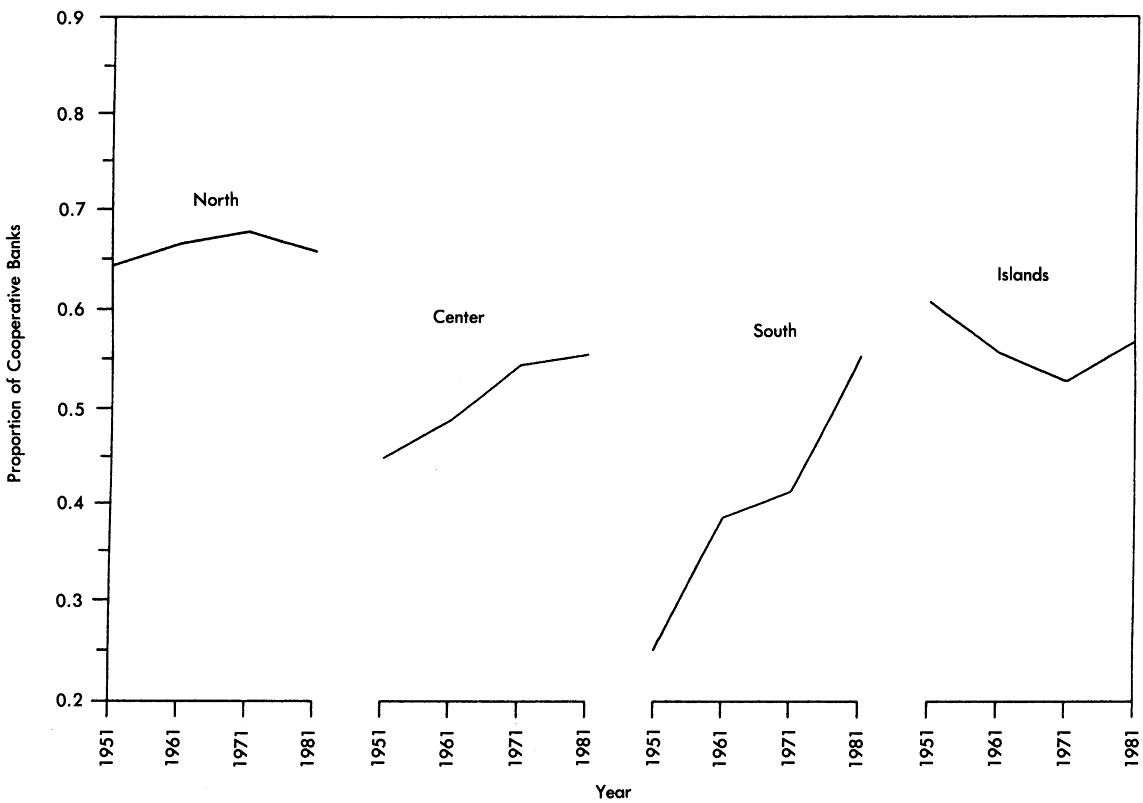


Figure 1 shows that at the national level, a vigorous founding activity increased RCBs as a proportion of the total number of banks from 59 percent in 1964 to 67 percent in 1989. Figure 2 shows that this percentage varied significantly across geographical macroareas of the country (north, south, center, and islands). There were 768 RCBs in existence at the end of 1964. This figure remained approximately stable until 1967–68, when monetary authorities enforced a general block on entry into the banking industry. The block lasted until 1971–72, but the population of RCBs continued to decline until 1976, when it reached its minimum at 641 units. During the 1970s and '80s, density gradually increased to reach 726 units in 1989, its maximum level in 20 years. These aggregate fluctuations in population density at the national level mask the heterogeneity of local responses to general ecological and institutional processes. Figure 3 shows that in Sicily (I1) and in the southern regions (S1–S5), population density increased, while it stayed virtually unchanged in the central regions (C1–C3) and declined steeply in the northern regions (N1–N4), where initial density was higher and the population was established earlier.

During the period 1964–1989, the size of the average rural cooperative bank, measured in terms of number of

Figure 2. Proportion of cooperative banks by geographical macro-area at ten-year intervals, 1951–1981.



branches, was less than two, reflecting the legal prohibition on branching imposed on rural cooperative banks, which remained essentially unit banks everywhere. At the end of 1988 RCBs accounted for 9 percent of the total number of bank branches in the country (12 percent in the sample). During the study period, the average size of the eight core national banks, all state-owned or state-controlled, grew from 250 to 387 (in terms of number of branches). At the end of 1988 the core national banks jointly accounted for about 22 percent of the total number of branches in the Italian banking industry (21 percent in the sample).

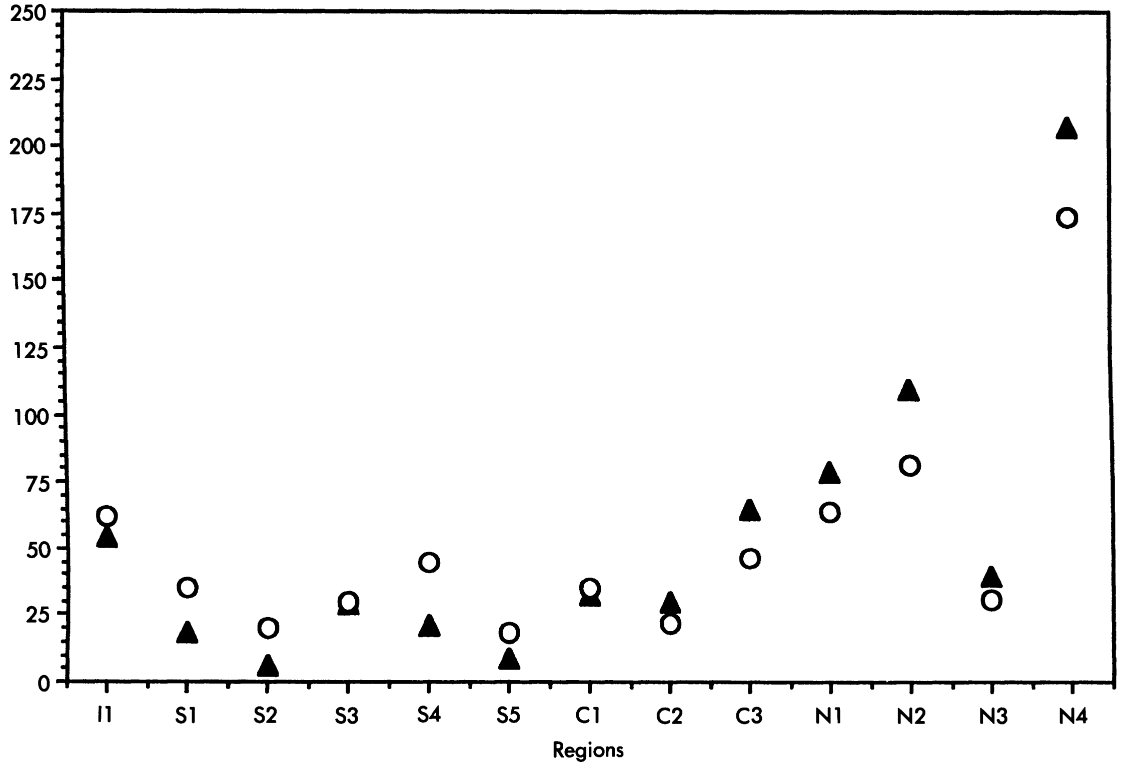
Empirical Specifications

The models of organizational founding rates estimated in this study are based on the following specification:

$$\lambda_{it}(\beta_1, \beta_2, \gamma, \pi, \theta_j | \alpha_i) = \exp\left(\alpha_i + \beta_1 N_{it} + \beta_2 N_{it}^2 + \gamma P_1 + \pi Q_{it} + \sum_{j=1}^{m} \theta_j X_{itj}\right), \quad (1)$$

where λ_{it} is the organizational founding rate in region i in time period t, i = 1, 2, . . . , N indexes the regions in the sample, t = 1, 2, . . . , T identifies the time periods for which data at the regional level are available (1964–1988), and j = 1, 2, . . . , m indexes the m-dimensional vector of measured control factors.

Figure 3. Variations in regional density of cooperative banks during the study period.*



* Black triangles along the Y-axis indicate organizational density in 1964, and white circles indicate organizational density in 1988. On the X-axis, which reports the different regions in the sample, I = island, S = South, C = Center, N = North.

The variables contained in equation (1) included N_{it}, the intrapopulation density of rural cooperative banks measured at the regional level, and its square (N_{it}^2); P_1, a period effect corresponding to the temporary legal block to entry (1968–1971) into the Italian banking industry imposed by the central credit authorities; and Q_{it}, the share of the total branches jointly controlled by the core national banks operating in Italy in the ith region. Finally, α_i is the heterogeneity component for the ith region. As discussed in detail below, α_i can be considered as a fixed region-specific effect, as a random draw from some underlying probability distribution, or it can be estimated directly from the data.

According to the density dependence hypothesis, I would expect a positive first-order effect ($\beta_1 > 0$) and a negative second order effect ($\beta_2 < 0$) of density. Hannan and Carroll (1992) insisted that failure to incorporate data on the population's early history may obscure the legitimating effects of population density; hence, the possibility of strictly competitive density effects cannot be ruled out in the present case. Although the current sample may be of limited value for testing the theory of density-dependent founding rates directly, the empirical results reported below suggest that it is unlikely that left-truncation has generated aberrant findings, because the estimated turning point in the relation between density and the founding rate (i.e., the estimated maximum rate) always falls within the observed range of density.

I expect a significantly negative effect of P_1 ($\gamma < 0$), signalling the effectiveness of the legal block to entry into the industry enforced by the central bank between 1968 and 1971. No other regulatory event significantly affected organizational founding rates in the population during the observation period. This was not surprising given the relatively short period of time covered by the study and the stability of the Italian banking legislation during the last 50 years (Costi, 1986; Lomi, 1995).

The share of productive capacity controlled by the core banks in each region (measured by the proportion of branches that they jointly control), Q_{it}, is included to control for resource partitioning, which a previous study found to affect organizational creation in the Italian banking industry significantly (Freeman and Lomi, 1994). The number of branches banking organizations control is proposed as an appropriate proxy for the concept of "productive capacity installed," which, according to Winter (1990), captures the differential burden placed by large organizations on the environmental carrying capacity and better represents the intensity of competition in an industry (Freeman and Lomi, 1994). According to the resource partitioning hypothesis, competition among large generalist organizations to occupy the center of the market will free resources at the periphery of the system that are most likely to be absorbed by specialist organizations (Carroll, 1985). In concentrated markets with few large generalist organizations, specialists may be able to exploit more resources without engaging in or at least before engaging in direct competition with larger generalist organizations. The process of resource partitioning, according to which large organizations can exploit the advantages of generalism by continuing to grow and small organizations exploit the advantages of specialism by not growing at all, can also be invoked as a rationale for the dramatic organizational size differentials resulting in the familiar skewed size distribution characterizing many business sectors (Ijiri and Simon, 1977; Hannan, Ranger-Moore, and Banaszak-Holl, 1990), including the Italian banking industry (Landi, 1990). Hence, Q_{it} is expected to have a positive impact on the founding rate, indicating the presence of processes of resource partitioning in the Italian banking industry (prediction: $\pi > 0$).

Environmental covariates. The following environmental covariates contained in the term $\sum_{j=1}^{m} \theta_j X_{itj}$ of equation (1) are included to control for observable differences in general economic, social, and competitive conditions across Italian geographical regions that may affect the organizational founding rate.

Agricultural employment (X_1). Rural cooperative banks are credit institutions specializing in financial and banking services to agriculture and small businesses. The number of agricultural workers in each region is used as a proxy for agricultural intensity and is predicted to have a positive effect on the founding rate of rural cooperative banks, since regions in which agricultural activity is prevalent are likely to provide a supportive environment for these specialist financial intermediaries, which render financial services to agricultural and craft businesses (prediction: $\theta_1 > 0$). Other variables were used in the early stage of the research to control for market and community size, including regional population, labor force, and employment, but none of these seemed to make a difference in terms of the conclusions supported by the model presented.

Number of branches of popular cooperative banks existing in the region (X_2). Popular cooperative banks (PCBs) are the second kind of cooperative bank existing in Italy. Unlike RCBs, which are regulated as cooperative organizations, PCBs are regulated as banks and their status of cooperative organizations is much more ambiguous in the current legislation (Costi, 1986). The number of branches is introduced as a proxy for the development of credit cooperation in the region. I offer no hypothesis about the direction of interpopulation interdependence, which can be either negative, signalling the existence of competition between the two subpopulations of cooperative banking organizations, or positive, indicating community-level mutualism. A previous analysis of the period 1936–1989 conducted at the national level revealed patterns of asymmetric mutualism between the two subpopulations of cooperative banks (Lomi, 1995) (prediction: $\theta_2 \neq 0$).

Total number of cooperative organizations in the macro-area (X_3). This variable represents the total number of cooperative organizations (consumer, producer, construction, transportation, and agricultural) present in the four standard geographical macro-areas into which Italy is usually partitioned for statistical purposes (north, center, south, and islands). This variable is assumed to capture the diffusion of the cooperative movement in different areas of the country. The density of cooperative organizations is expected to have a positive effect on the founding rate of RCBs, reflecting both institutional processes of mimicry at the time of founding, according to which widespread experience of a given organizational form accelerates the founding rate of organizations assuming that form (DiMaggio and Powell, 1983; Freeman, 1990), as well as community-level mutualism (both commensalistic and symbiotic) among populations of organizations with the same form (i.e., structure of property rights). The presence of a strong cooperative sector beyond the strictly regional scope signals the legitimation of this kind of organization in the area and is likely to contribute to a supportive institutional environment for the founding of cooperative banks (prediction: $\theta_3 > 0$).

Total value of bank deposits in the region (X_4). As the main resource of banking organizations, the value of deposits collected by the entire banking system in the region is introduced to control for the effects of local differences in the carrying capacity of the environment for the population. Following the specialized economic literature on banking (Conigliani, 1983), deposits are used here as a proxy for demand for banking services and are expected to affect the founding of cooperative banks positively (prediction: $\theta_4 > 0$).

Total number of mergers and acquisitions among RCBs (X_5). Based on the resource partitioning argument, intrapopulation mergers and acquisitions among RCBs is included to control for organizational growth, since external growth is the only growth strategy legally available to these specialist unit banks that are legally barred from growing by branching (prediction: $\theta_5 > 0$).

Methods of Analysis

Organizational foundings can be considered as an instance of an arrival process (Hannan and Freeman, 1989: chap. 8; Hannan, 1991). The parameter of interest in this process is the arrival rate, defined as the instantaneous probability of arriving at state $(y + 1)$ at time $(t + \Delta t)$, as given in the following:

$$\lambda_y(t) = \lim_{\Delta t \downarrow 0} \frac{\Pr\{Y(t + \Delta t) - Y(t) = 1 \mid Y(t) = y\}}{\Delta t} \tag{2}$$

where $Y(t)$ is the cumulative number of foundings up to time t. The baseline model formulation assumes that $\lambda_y(t) = \lambda$ and that the (conditional) probability of y_t arrivals in any time interval is governed by the probability law:

$$\Pr(Y_t = y_t x_t) = \frac{[e^{-\lambda(x_t)}\lambda(x_t)^{y_t}]}{y_t!}, \tag{3}$$

where the expected number of foundings in each period $E(Y_t) = \lambda_t$ equals the variance. Given that most available data on founding, including those I used, are available in the form of yearly counts—a discrete but noncategorical variable—it is appropriate to use the Poisson regression model to estimate covariate effects on founding probabilities (Hausman, Hall, and Griliches, 1984; Maddala, 1989: chap. 2; Barron, 1992b).

In pooled cross-sectional time series data, where repeated observations on individual units are available over time, heterogeneity in the arrival rate can be captured by a set of region-specific dummy variables. This fixed-effects approach has the advantage of making no assumptions about the distribution of heterogeneity. The disadvantages are that (1) the number of parameters to estimate grows with the sample size (Chamberlain, 1985), (2) parameters of time-invariant covariates cannot be easily estimated (Judge et al., 1985; Reader, 1993), and (3) the estimates cannot be used to predict founding rates outside the estimation sample (Hsiao, 1986). These potentially serious problems with the fixed-effects approach can be overcome by allowing the Poisson parameter λ to vary randomly across individual units according to a certain probability distribution (Hausman, Hall, and Griliches, 1984).

Following this method of representing heterogeneity in a stochastic model by a mixing distribution used to compound the baseline Poisson process, previous research on the ecology of organizational founding has typically relied on the negative binomial specification derived from the baseline Poisson model with gamma mixing (Barron and Hannan, 1991; Hannan, 1991; Barron, 1992b; Hannan and Carroll, 1992; Ranger-Moore, Banaszak-Holl, and Hannan, 1991).[2] A potential problem with this representation of heterogeneity by a mixing distribution is that there is little theoretical

2
This model corresponds to the "apparent contagion" model used in biometrics, according to which units at risk have a constant but unequal probability of experiencing an event (Cameron and Trivedi, 1986).

reason for choosing one specific parametric form from among a number of equally plausible others (Sichel, 1982; Dunn, Reader, and Wrigley, 1987; Reader, 1993). Another problem is that an inappropriate functional specification for heterogeneity will produce biased estimates of the parameters included in the model (Heckman and Singer, 1982, 1984; Davies and Crouchley, 1985; Yamaguchi, 1986). Hence the choice of form for the distribution of unobservables is important because alternative distributional assumptions have different qualitative implications in terms of the underlying theoretical propositions and thus they imply different models (Cameron and Trivedi, 1986; Hannan and Carroll, 1992). Unfortunately, founding processes turn out to be sensitive to distributional assumptions, and no single parametric model was found to be uniformly best, in a statistical sense, in all populations.[3] On the basis of their extensive empirical experience, Hannan and Carroll (1992: chap. 4) concluded that it is premature to settle for a specific parametric model but that it is unlikely that theory will ever be able to indicate a distribution of unobservables.

With many possible alternatives and no theoretical reason for preferring one mixing distribution over another, it is essential to consider the possible failure of any one specific parametric form to represent adequately the distribution of the unobservables. To minimize the impact of arbitrary distributional assumptions on inferential results, Heckman and Singer (1984) proposed a semiparametric random-effects approach in the context of single-spell Weibull duration models with two states, and Brännäs and Rosenqvist (1994) recently extended this procedure to count data models. One limitation of this semiparametric random-effects Poisson model is that it assumes that heterogeneity arises only from differences in the mean baseline rates, while the coefficients are considered fixed across sample units. This model is inappropriate if the population is heterogeneous also with respect to the impact of some of the explanatory variables. Given that the central theoretical issue addressed by the paper is the heterogeneous local responses of the founding rates due to location dependence, this assumption is relaxed to allow for heterogeneity both in the mean baseline founding rate (the intercepts) and in how the founding rates respond to general population-level processes related to density and resource partitioning (the coefficients). This modified version of the semiparametric random-effects Poisson model deals with unobserved heterogeneity in organizational founding rates in two ways: First, the mean event rate is allowed to vary across a finite number of unobserved classes, defined in terms of a discrete mixture distribution estimated from the data. Second, the mean event rate is allowed to vary within classes depending on the value of the explanatory variables. The main limitation of this discrete mixture Poisson model is that the observations in each latent class are assumed to be generated by a homogeneous Poisson process. The implication of this assumption is that if additional heterogeneity is present within each class over and above that captured by the measured covariates, the estimated standard errors of the parameters would be biased downward.

3
Hannan and Carroll (1992: 92–93) found that generalized Yule models fit banks and brewers well, but newspaper populations were best described by a log-quadratic approximation, and founding rates in the populations of labor unions and insurance companies were best represented by a Gompertz specification. They also found that the parameter estimates in the brewer population were sensitive to alternative distributional assumptions.

Semiparametric random-effects Poisson model. The main advantage of the semiparametric maximum likelihood (ML) approach developed by Heckman and Singer (1984) is the introduction of flexibility in the shape of the mixing distribution by assuming that the density function of the heterogeneity components can be approximated by a finite number of discrete support points (Trussell and Richards, 1985). This estimation procedure involves determining the location of these points and their associated probability masses empirically from the data, i.e., without specifying any probability distribution for the unobservable heterogeneity terms (Brännäs and Rosenqvist, 1994; Gupta and Chintagunta, 1994). The analysis of the pooled cross-sectional time series data of organizational founding is based on the following Poisson probability specification:

$$\Pr(Y_{it} = y_{it}) = \frac{e^{-\lambda_{it}}\lambda_{it}^{y_{it}}}{y_{it}!} \tag{4}$$

with intensity parameter

$$\lambda_{it} = \exp\left(\alpha + \sum_{j=1}^{p} \beta_j x_{ijt}\right), \tag{5}$$

where x_{ijt} is the value of independent variable x_j for region i at time t. The term α in equation (5) brings about overdispersion in y_{it}, the aggregate series of organizational foundings.[4] Because of differences due to unobservable (unmeasured) characteristics related to local conditions, the founding rates may vary across regions. To account for this, it is assumed that a region's intrinsic founding rate, represented by the intercept term α, varies across regions. Specifically, if α is a realization from some unknown underlying probability distribution $f(\alpha)$, then the (unconditional) likelihood function for the ith region is:

$$L_{i|\alpha} = \int_{\alpha} \prod_{t=1}^{T} \exp(-\lambda_{it}) \frac{(-\lambda_{it})^{y_{it}}}{y_{it}!} f(\alpha)d\alpha. \tag{6}$$

Situations in which $f(\alpha)$ is assumed to have a known parametric form generate the standard random-effects Poisson model, which is based on the assumption that $f(\alpha) \sim \Gamma$ across cross-sectional units (Hausman, Hall, and Griliches, 1984).

Ideally, distributional assumptions should reflect some theory or prior information about the process generating the observations but, in practice, distributional assumptions just reflect computational convenience and familiarity (Reader, 1993). Lacking specific theoretical indications, it is important to reduce the impact on the inferential results of arbitrary assumptions about the distribution of unobservables (Heckman and Singer, 1984; Trussell and Richards, 1985). To accomplish this, the semiparametric random-effects Poisson model estimated below does not assume any parametric form for $f(\alpha)$, which is approximated by a finite number of support points estimated directly from the data. The number of support points is determined by an iterative procedure that adds points until the inclusion of an additional point fails to improve the likelihood of the model significantly.

4
This is readily proved. Let $\lambda_{it} = \exp(\beta'X)$, and $\alpha = \exp(\theta)$ so that $\lambda\alpha = \exp(\theta + \beta'X)$. Assume that α is a random draw from a probability distribution and (for identifiability) that $E[\alpha] = 1$, and $\mathrm{Var}[\alpha] = \sigma^2$. Then

$$\begin{aligned}\mathrm{Var}(y_{it}) &= E[\mathrm{Var}(y_{it}|\alpha)] + \mathrm{Var}[E(y_{it}|\alpha)]\\ &= E(\lambda_{it}\alpha) + \mathrm{Var}(\lambda_{it}\alpha)\\ &= \lambda_{it}E(\alpha) + \lambda_{it}^2\mathrm{Var}(\alpha)\\ &= \lambda_{it}(1 + \sigma^2\lambda_{it}),\end{aligned}$$

so that the variance to mean ratio is $(1 + \sigma^2\lambda_{it}) > 1$. For this reason, the heterogeneous Poisson model estimated here is called the overdispersed Poisson model (Brännäs and Rosenqvist, 1994).

The support points (α_k) can be interpreted as the unconditional probability that a unit belongs to class k. It follows that for each class of the discrete mixing distribution the following constraints have to be satisfied:

$$\sum_{k=1}^{K} \alpha_k = 1, \text{ and } 0 < \alpha_k < 1. \tag{7}$$

Conditional on the kth support point, the semiparametric random-effects formulation of the Poisson parameter is:

$$\lambda_{it}(\beta | \alpha_k) = \exp\left(\alpha_k + \sum_{j=1}^{p} \beta_j x_{ijt}\right), \tag{8}$$

where α_k ($k = 1, 2, \ldots, K$) is the heterogeneity compound, and β is a p-dimensional vector of unknown parameters to be estimated. If the discrete distribution of α has K support points ($\alpha_1, \alpha_2, \ldots, \alpha_k$) with K associated probabilities $\{\Pr(\alpha_1), \Pr(\alpha_2), \ldots, \Pr(\alpha_k)\}$, then the contribution of each cross-sectional unit (region) to the likelihood function for the entire sample is:

$$l_i = \sum_{k=1}^{K} \prod_{t=1}^{T} \exp[-\lambda_{it}(\beta | \alpha)] \frac{[\lambda_{it}(\beta | \alpha_k)]^{y_{it}}}{y_{it}!} \Pr(\alpha_k). \tag{9}$$

The likelihood function for the entire sample is therefore:

$$L = \prod_{i=1}^{N} L_i. \tag{10}$$

Detailed discussions about the numerical algorithms that can be used to maximize L over the parameters can be found in Brännäs (1992), Brännäs and Rosenqvist (1994), and Wedel et al. (1993).

Model selection criteria. A potential problem with the semiparametric specification of the likelihood function is the formal comparison among different models, since models with different numbers of support points are based on different specifications of the heterogeneity term and are therefore non-nested. To assess the consistency of alternative non-nested model specifications with the data, and to take into account the larger number of parameters to be estimated in the heterogeneous random-effects Poisson model, the models reported in the following section are compared on the basis of the $\bar{\rho}^2$ statistic (Horowitz, 1983; Ben-Akiva and Lerman, 1985) and the shortest data description criterion (Rissanen, 1978).

The $\bar{\rho}^2$ test for model specification is a likelihood ratio test adjusted to account for differences in degrees of freedom across non-nested models and is defined as

$$\bar{\rho}^2 = 1 - \frac{L_f - \eta_f}{L(0)}, \tag{11}$$

where L_f is the log-likelihood of the full model, η_f is the number of parameters, and L(0) is the log-likelihood of the restricted model containing a constant term only. The model with the highest $\bar{\rho}^2$ value will be the one most consistent with the data.

The shortest data description criterion (SDDC), takes the form of a likelihood function modified to account for overparametrization. The SDDC used here—sometimes referred to as Rissanen's criterion (Rissanen, 1978, 1985)—is defined as SDDC $= -2\log[\max L(k)] + \phi(k,n)$, where $[\max L(k)]$ denotes the maximum value of the likelihood over the parameters, and $\phi(k,n)$ is a logarithmic function of the number of independent parameters to be estimated (k) and the number of observations (n) and represents the cost of fitting an additional parameter. The SDDC, which was originally derived as a solution to the problem of minimum-bit representation of a signal (Rissanen, 1978), has been shown to outperform similar model-selection criteria both empirically and theoretically (Sclove, 1987). A model is good if it gives a small value of the SDDC relative to the value given by competing models.

The empirical analysis deals with three questions: (1) Do founding rates of cooperative banks vary systematically across levels of spatial aggregation, and, if so, what level of analysis is most consistent with the theoretical predictions of ecological theories? (2) Are models accounting for unobserved heterogeneity due to location dependence significantly more consistent with the data than corresponding models that ignore unobserved heterogeneity? and (3) What are the implications of location dependence for organizational founding rates?

I answer these questions in three steps. First, I compare different estimates of the baseline model with local (regional) and nonlocal (national) density specifications. Second, I introduce unobserved heterogeneity by allowing the mean event rate to vary across a finite number of unobservable classes defined in terms of a discrete mixture distribution estimated from the data. Finally, I use the best-model specification from the second step and allow the mean event rate to vary within the heterogeneous classes, depending on the value of the explanatory variables of theoretical interest.

RESULTS

Table 1 reports ML estimates of the baseline models. The models differ in how they specify the form of density dependence and the level of aggregation at which density is measured. Column 1 contains the estimates of the parameter of interest without controlling for density effects. Columns 2 and 3 contain the linear and nonlinear effects of density specified at a local (regional) level, while in the model estimated in columns 4 and 5, density is specified at the national level. According to these estimates, the model accounting for the nonlinear effects of regional density (column 3) significantly improves on models that do not control for density (column 1). The model based on the linear density specification (column 2) is not statistically different from the model in which density is omitted altogether (column 1). Allowing the effects of regional density to be nonmonotonic increases the fit significantly.

Table 1

Poisson Regression Models: ML Estimates of Equation (1)*

Variable	1	2	3	4	5
Intercept	−2.261••	−2.238••	−2.548•	−4.17••	−11.486
	(.384)	(.391)	(.417)	(1.112)	(5.982)
National density				.0028	.026
				(.0016)	(.019)
(National density)^2/100					.0019
					(.0015)
Regional density		.0007	.027•		
		(.0023)	(.012)		
(Regional density)^2/100			−.013•		
			(.006)		
Core banks' share	3.198••	3.155••	3.574••	3.180••	3.071••
	(.671)	(.684)	(.701)	(.671)	(.675)
P1 (1968–1971)	−1.938••	−1.941••	−2.017••	−2.025••	−2.095••
	(.459)	(.459)	(.461)	(.462)	(.467)
Agricultural employment	.012	.012	.003	.001	.001
	(.008)	(.008)	(.009)	(.008)	(.008)
Agencies of PCBs	.004	.005	−.009	.007	.005
	(.006)	(.006)	(.008)	(.006)	(.006)
Cooperative organizations/1,000	.300••	.307••	.250••	.325••	.283••
	(.130)	(.115)	(.100)	(.129)	(.133)
Deposits/10,000	.058	.058	.030	.048	.051
	(.059)	(.059)	(.060)	(.058)	(.059)
RCB's M&A	.236••	.244••	.251••	.270••	.294••
	(.036)	(.035)	(.0363)	(.0328)	(.039)
Log likelihood	−310.99	−310.94	−308.25	−309.34	−308.50
Chi-square	392.87	392.22	385.95	390.19	386.06
G^2	339.51	339.42	334.04	366.22	334.54
Degrees of freedom	8	9	10	10	10
Cases	325	325	325	325	325

• $p < 0.05$; •• $p < 0.01$.
* Asymptotic standard errors are in parentheses.

The results differ when density is specified at the national level. The estimates reported in columns 4 and 5 show no significant effect of national density on organizational founding rates. Comparing columns 3 and 5 in Table 1 shows that the ratio between the regional and national first-order effects of density is close to unity (1.04), indicating virtually no difference in legitimation across models based on local and nonlocal specifications of density. The ratio between the regional and national second-order effects of density on the founding rates shows that competition is about seven times (6.84) stronger at the regional than at the national level. On the basis of these simple considerations, the analysis provides solid support for Hannan and Carroll's (1992) conjecture that differences in strength between estimates based on local versus nonlocal specifications of density are greater for the second-order effect, which is associated with density-dependent competition, than for the first-order effect, which is associated with density-dependent legitimation.

The estimates reported in Table 1 also provide strong support for the resource partitioning hypothesis discussed above. Concentration in the banking industry significantly increased the founding rates of specialist cooperative banks. Similarly, concentration in the cooperative banking sector due to mergers favored the appearance of new unit banks. The other control factors affected the founding rate in the

Table 2

Parametric and Semiparametric Random Effects ML Estimates of Equation (1)*

Variable	No heterogeneity	Negative binomial	2-point heterogeneity	3-point heterogeneity
Intercept 1 (α_1)	−2.548•	−2.498••	−3.196••	−2.513••
	(.417)	(.501)	(.509)	(.586)
Intercept 2 (α_2)			−4.356••	−3.323••
			(.622)	(.711)
Intercept 3 (α_3)				−4.080••
				(.681)
Regional density	.027•	.027•	.021	.007
	(.012)	(.013)	(.013)	(.017)
(Regional density)^2/100	−.013•	−.014•	−.010	−.025
	(.006)	(.007)	(.0624)	(.082)
Core banks' share	3.574••	3.597••	3.825••	2.435••
	(.701)	(.983)	(.718)	(.927)
P1 (1968–1971)	−2.017••	−2.054••	−1.762••	−1.760••
	(.461)	(.518)	(.464)	(.439)
Agricultural employment	.003	.002	.026•	.0316••
	(.009)	(.010)	(.011)	(.0135)
Agencies of PCBs	−.009	−.009	−.008	−.004
	(.008)	(.012)	(.009)	(.012)
Cooperative organizations/1,000	.250••	.230	.679••	.059••
	(.100)	(.154)	(.159)	(.017)
Deposits/10,000	.030	.033	.0629	.129
	(.060)	(.075)	(.060)	(.072)
RCB's M&A	.251••	.263••	.290••	.287••
	(.036)	(.068)	(.039)	(.035)
Overdispersion parameter		.236		
		(.151)		
Log likelihood	−308.25	−306.71	−298.98	−297.90
$\bar{\rho}^2$	.144	.145	.164	.161
SDDC	647.41	650.20	634.47	635.98
Degrees of freedom	10	11	12	14
Cases	325	325	325	325

•$p < .05$; ••$p < .01$.

* Estimated standard errors are in parentheses. For ease of comparison between the baseline model and the heterogeneous semiparametric Poisson model, column 1 contains the same estimates reported in column 3 of Table 1.

direction predicted. The founding rate dropped almost to zero as a consequence of legal restrictions imposed by central credit authorities. The presence of cooperative organizations in other economic sectors increased the founding rate of cooperative banks. The effects of agricultural intensity on the founding rate of the region and deposits turned out to be statistically weak but in the direction predicted. Finally, changes in size in the subpopulation of popular cooperative banks had no significant effect on the founding rate of rural cooperative banks, indicating that the niches occupied by the two populations of mutual banks did not overlap.

The second question concerns the representation of unobserved heterogeneity due to location dependence. The results, accounting for unobserved heterogeneity without imposing any parametric restrictions, are reported in Table 2. The semiparametric model accounting for unobserved heterogeneity in column 3 is clearly more consistent with the data than the corresponding baseline model (column 1, Table 2) estimated without correcting for unobserved heterogeneity. Formally, the computed $\bar{\rho}^2$ value for the model without heterogeneity containing the nonlinear

Table 3

Semiparametric Random Effects ML Estimates of Equation (7)*

Variable	2-point solution	3-point solution
$Intercept_1$	−2.307**	−1.960
	(.432)	(1.120)
$Density_1$	.008	−.068*
	(.020)	(.003)
$(Density_1)^2/100$	−.003	.003*
	(.080)	(.001)
Core banks' $share_1$	3.017**	.127
	(.843)	(.325)
$Intercept_2$	−8.247**	−7.108**
	(1.901)	(2.987)
$Density_2$	.016†	0.068
	(.008)	(.008)
$(Density_2)^2/100$	−.010	−.020
	(.010)	(.050)
Core banks' $share_2$	5.901**	3.054
	(2.527)	(3.756)
$Intercept_3$	–	−2.045**
		(.612)
$Density_3$	–	.003
		(.004)
$(Density_3)^2/100$	–	−.020
		(.057)
Core banks' $share_3$	–	−2.602
		(1.379)
P1 (1968–1971)	−1.886**	−1.807**
	(.473)	(.467)
Agricultural employment	.002	.006**
	(.001)	(.002)
Agencies of PCBs	−.001	−.003
	(.001)	(.009)
Cooperative organizations/1,000	.041*	.078**
	(.018)	(.021)
Deposits/10,000	.002	.002*
	(.001)	(.0008)
RCB's M&A	0.264**	0.031**
	(.0037)	(.004)
Log likelihood	−299.03	−291.10
$\bar{\rho}^2$	.155	.163
SDDC	642.79	640.52
Degrees of freedom	15	20
Cases	325	325

• $p < .05$; •• $p < .01$.
* Estimated standard errors are in parentheses.
† Significant at $p < .10$.

density effects specified at the regional level (column 3, Table 1) is $\bar{\rho}^2 = 0.1439$; for the corresponding heterogeneous model (column 2, Table 2) $\bar{\rho}_H^2 = 0.1635$. The test statistic for the formal comparisons is $\Pr(\bar{\rho}_H^2 - \bar{\rho}^2 > \tau) \leq \Phi\{-[-2\tau L(0) + (\eta_H - \eta)]^{1/2}\}$, where $\tau \leq 0$, Φ denotes the cumulative distribution function of the normal distribution, and η_H, η are the numbers of parameters for the two models (Ben-Akiva and Lerman, 1985). Substituting the numerical values in the expression of the test statistics gives $\Pr(\bar{\rho}_H^2 - \bar{\rho}^2 > \tau) \leq \Phi\{-4.19206\}$, which indicates that the probability that this difference could have occurred by chance is less than 0.001. A similar indication comes from comparing the two models on the basis of the SDDC. In terms of these model-selection criteria, the semiparametric random-effects model also outperforms its parametric rival, reported in the second column of Table 2. The negative binomial model fails to improve significantly over the baseline Poisson model and is unable to detect the presence of overdispersion in the data. The SDDC clearly indicates that a negative binomial process is not what produced these observations.

A related problem in model specification and selection concerns the number of support points necessary to approximate the distribution of the heterogeneity component. Both the SDDC and the $\bar{\rho}^2$ test for model specification indicate that the semiparametric model estimated with a three-point heterogeneity component fails to improve over the corresponding two-point model, and in fact does slightly worse in terms of overall fit and proves less informative. Going from a two- to a three-mass-points representation of the heterogeneity in the sample changes the likelihood only marginally, reflecting the lack of improvement in the fit of the model to the data. The conclusion is that heterogeneity in the sample can be described semiparametrically by a two-point distribution. In general, these estimates indicate that models accounting for unobserved heterogeneity are more consistent with the data than the simple Poisson regression model after adjusting for the greater number of parameters to be estimated. The same basic conclusion holds when models based on negative binomial assumptions about the distribution of heterogeneity components are estimated.

Figure 4. Multiplier of the founding rate based on the estimates of models 1 and 3 in Table 2. (Vertical dotted lines mark the observed range of density.)

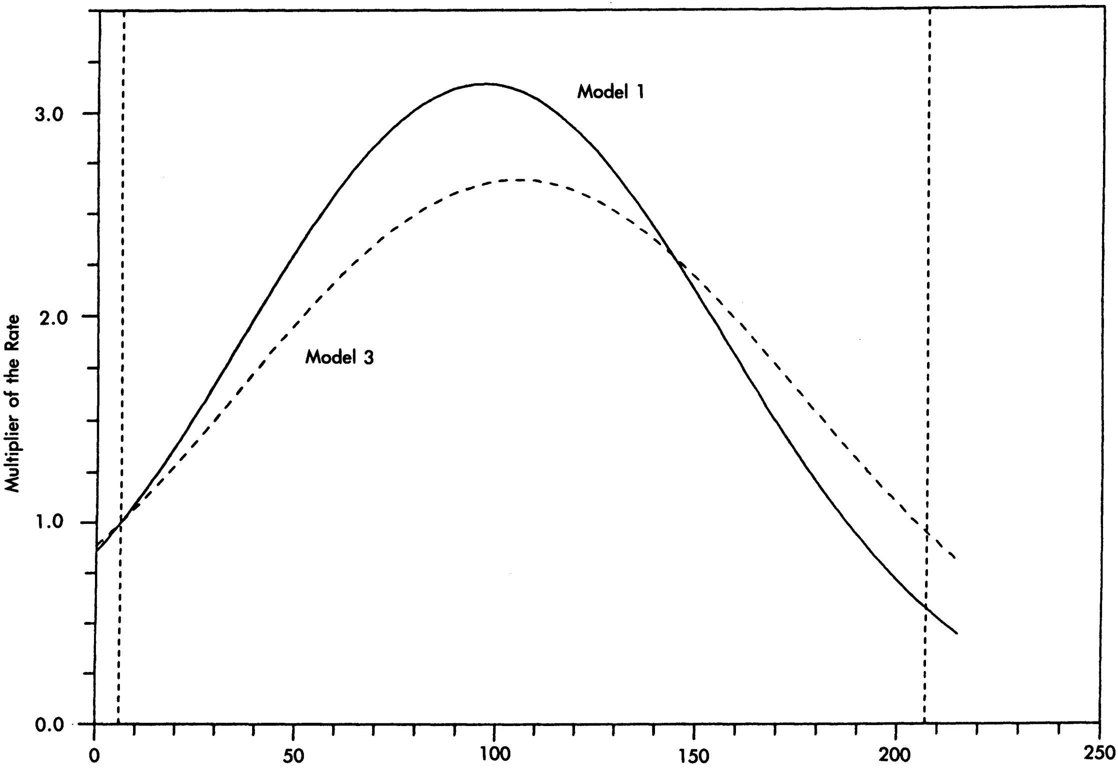
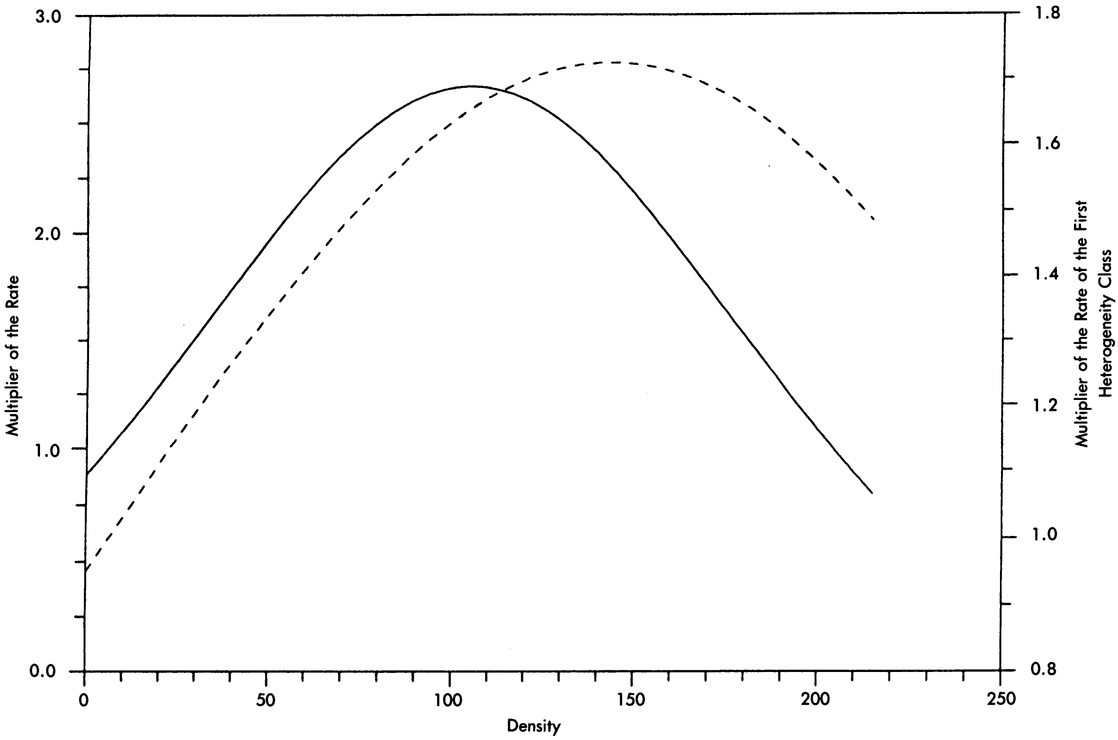


Figure 5a. Multiplier of the founding rate in the first segment (dashed line = estimates of model 1 in Table 3) and in the population, allowing for heterogeneity between segments only (solid line = estimates of model 3 in Table 2).



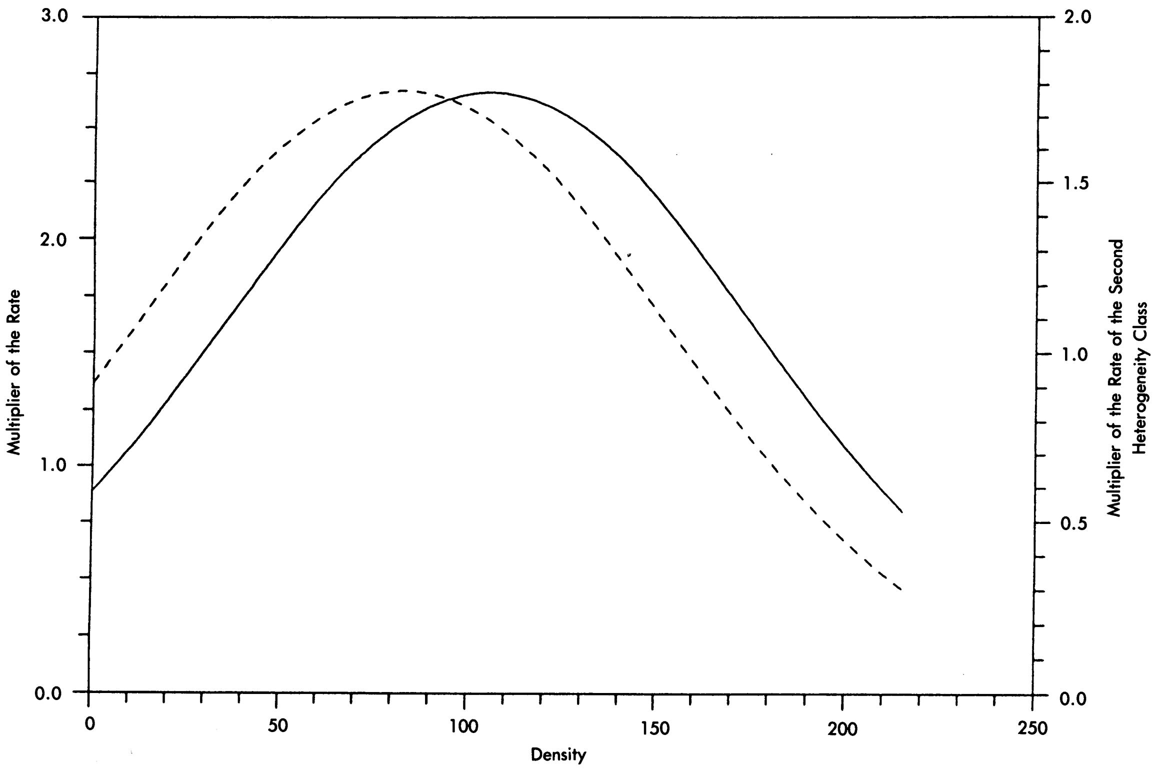
The third question concerns the specific effects of heterogeneity on organizational founding rates. Column 3 in Table 2 shows that the nonlinear effects of regional density maintain the theoretically predicted signs and magnitudes but lose statistical significance when unobserved heterogeneity is accounted for without imposing any parametric assumption. Table 3 reports the estimates produced by allowing the mean event rate to vary within the heterogeneous classes identified in terms of the support points, depending on the value of the explanatory variables of theoretical interest. The effects of density and concentration on the founding rates are now allowed to vary across classes to test for heterogeneity in local responses to global population processes related to density and resource partitioning.

The density effects are in the theoretically predicted direction but not statistically significant. In terms of magnitude, the effects of density are much stronger in the second segment of the population than in the first. The values of the control variables remain numerically stable and in the direction expected. When three heterogeneous classes are assumed, the estimates show clear signs of misspecification. As before, tests for model specification strongly favor the two-point model.

Qualitative Implications

The implications of these findings for the population ecology of cooperative banks can be seen in Figure 4, which plots

Figure 5b. Multiplier of the founding rate in the second segment (dashed line = estimates of model 1 in Table 3) and in the population, allowing for heterogeneity between segments only (solid line = estimates of model 3 in Table 2).



the estimated relationship between the number of banks and the rate implied by the estimates of the statistically best model, controlling for unobserved heterogeneity (Table 2, column 3), and by the baseline model (Table 2, column 1). The dotted vertical lines in the figure mark the minimum and maximum observed density in the sample. The multiplier of the rate implied by the heterogeneous model is unity at the minimum observed density (N_{min} = 6), and 0.94 at the maximum observed density (N_{max} = 207). At its maximum, when the density is 105 [$N(\lambda_{max})$], the rate (λ_{max}) is 2.7 times larger than the rate when the density is at its minimum and approximately three times larger than the (estimated) rate at zero density. According to the baseline model, the multiplier of the rate at minimum density [$\lambda(N_{min})$] is unity, but it is only 0.57 at maximum density. The founding rate reaches its maximum of 3.14 when the density is 97. These figures indicate that models neglecting heterogeneity tend to overestimate the effects of density on the founding rate. The rate rises faster with density and reaches a higher maximum earlier, at lower levels of density. The baseline model also implies that the founding rate drops off faster after reaching its peak than the rate as estimated by the heterogeneous model. A comparison of the multiplier of the rate at maximum density with the estimated maximum founding rate [$\lambda(N_{max})/\lambda_{max}$] for the two models indicates the strength of density-dependent competition. It shows how much the founding rate declines from its peak as density reaches its maximum value (Hannan and Carroll, 1992: chap. 4). At maximum density, the founding rate estimated by the baseline model is only 18 percent of the maximum rate, while in the heterogeneous model the rate at maximum sample density is 35 percent of the maximum rate.

Allowing the founding rate to vary within heterogeneous segments of the populations as a function of the covariates revealed the existence of two heterogeneous segments in the organizational population. Figure 5a plots (on different *y*-scales) the relationship between density and the arrival rate of organizations in the first heterogeneous segment (Table 3, column 1) and the same relationship according to the model estimated by controlling only for heterogeneity between segments (Table 2, column 3). The multiplier of the rate at the minimum observed density is again unity and reaches 1.5 at the maximum observed density. At its maximum, when density is 105, the rate is approximately 70 percent larger than the rate when density is at its (observed) minimum. In the first heterogeneous segment of the population, the effects of density dependence are weaker in that the rate rises more slowly and reaches a maximum at a higher level of density (142). At the maximum density, the arrival rate in the first heterogeneity class is still approximately 50 percent larger than the rate at the minimum density. The rate at maximum density is 90 percent of the maximum estimated rate, indicating a weak effect of density-dependent competition in this segment of the population.

The estimates in column 1 of Table 3 imply that the arrival rate of organizations in the second heterogeneous segment of the population is more sensitive to density than the rate in the first segment. The multiplier of the rate rises faster than average and reaches its maximum value of 1.8 when density is 82. After this point, the rate declines steeply. At maximum density, the rate is only 21 percent of the rate at its maximum and 37 percent of the rate at minimum density. As density increases beyond $N(\lambda_{max}) = 82$ and reaches its maximum, competition depresses the founding rates by approximately 80 percent. Organizations arriving into this second heterogeneous segment of the population will face a situation in which density-dependent competition dominates density-dependent legitimation at relatively low levels of density. Figure 5b plots (on different *y*-scales) the relationship between density and the arrival rates of organizations into the second heterogeneous segment of the population (Table 3, column 1) and the same relationship according to the model estimated by controlling only for heterogeneity between segments (Table 2, column 3) and assuming that heterogeneity in the population can be captured simply by letting intercepts be free to vary between segments.

Finally, the point of maximum founding rate in the models $[N(\lambda_{max}) = -\beta_1/\beta_2]$ is always within the observed range of density, so the function relating density to founding rates turns from positive to negative within the range of observations. Table 4 provides a summary of the qualitative implications of the estimates.

Table 4

Qualitative Implications of the Estimates of Density Dependence in Founding Rates*

Model	$\lambda(N_{min})$	$\lambda(N_{max})$	λ_{max}	$N(\lambda_{max})$	$\lambda(N_{max})/\lambda_{max}$
No heterogeneity	1	.567	3.142	97	.181
2-point heterogeneity	1	.941	2.665	105	.353
First segment	1	1.532	1.725	142	.888
Second segment	1	.373	1.782	82	.210

* Models estimated without accounting for heterogeneity are in the first row, models allowing the mean founding rate to vary between heterogeneous segments of the population are in the second row, and models allowing the founding rate to vary between and within two heterogeneous segments of the population are in the third and fourth rows. $\lambda(N_{min})$ and $\lambda(N_{max})$ denote the founding rate at the minimum and maximum values of density, respectively. λ_{max} denotes the founding rate at the maximum, when density is equal to $N(\lambda_{max})$. The last column shows the strength of density-dependent competition.

DISCUSSION AND CONCLUSIONS

Factors related to location dependence and heterogeneity are particularly troublesome in the study of organizational founding because it is not clear what selection forces operate upon, individual sources of organizational founding are not observable at the population level, and the set of units at risk is not well defined. Research on the ecology of organizational founding has partly circumvented these problems by considering the population as the unit experiencing the events. This approach is problematic, however, whenever identifiable segments of the population respond heterogeneously to general competitive and institutional processes. This study addressed these conceptual problems empirically by proposing a specific model of location dependence in organizational founding rates.

The analysis of founding rates in the population of Italian cooperative banks yields three main findings. First, the level of analysis is an important factor in specifying processes of organizational founding. Models of founding rates of cooperative banks are better specified at the regional level, while national density has no effect on founding rates. This conclusion is consistent with the results of Carroll and Wade (1991), who found that founding rates of American breweries are localized and that density dependence thus operates at lower levels of analysis. Empirical support was found also for Hannan and Carroll's (1992) conjecture that differences in strength between estimates based on local versus nonlocal specifications of density are greater for the second-order effect (associated with density-dependent competition) than for the first-order effect (associated with density-dependent legitimation). While there is no difference in legitimation across models based on local and nonlocal specifications of density, competition is about seven times stronger at the regional than at the national level. These results substantially qualify Zucker's (1989) intuition that different geographical areas differ systematically in terms of organizational processes of competition and legitimation. The extent to which these conclusions extend to more conventionally entrepreneurial organizations in less regulated sectors remains an important issue for future research in the ecology of organizations.

Second, models neglecting unobservable heterogeneity tend to overestimate the effects of density on founding rates. Analysis of the qualitative implications of the estimates revealed that the founding rate implied by the baseline model rises faster, reaches a higher maximum at a lower level of density, and, finally, drops off faster after reaching its peak than the rate as estimated by the heterogeneous model. As density reaches its observed maximum, the founding rate implied by the heterogeneous model declines from its peak to reach a value that is about twice as large as the value implied by the baseline model. At maximum density, the founding rate estimated by the heterogeneous model is virtually the same as the rate calculated at the observed minimum density, while the rate implied by the baseline model at maximum density is slightly more than 50 percent of the rate computed at minimum density. When heterogeneity is controlled for without making any assumption about the distribution of unobservable components across regions, the effects of density lose statistical significance but maintain a strong tendency to have the predicted nonmonotonic pattern.

Third, while the study supports theories that predict a nonmonotonic inverted-U shaped relationship between density and founding rates at the population level, the disaggregation of the overall population into distinct heterogeneous segments reveals a more complex relationship between location, density, and organizational founding rates. The analysis revealed the existence of two distinct heterogeneous segments in the population. When the mean rate is allowed to vary between segments, the effects of density become statistically weaker but maintain the predicted direction and magnitude. When the rate is allowed to vary within segments as a function of the covariates, evidence of heterogeneous response to general population processes is found. The ecological dynamics in the first segment of the population are comparable to that of Irish and San Francisco newspapers studied by Hannan and Carroll (1992) in that the relationship between density and founding rates is relatively flat from its peak to the point of maximum density. In this heterogeneous segment, the founding rate drops by approximately 10 percent from its peak as density reaches its observed maximum. The second heterogeneous segment of the population is characterized by a stronger competition from numbers, since the founding rate drops by 88 percent from its peak. The ecological dynamics in this second segment are comparable to that of Manhattan banks studied by Hannan and Carroll (1992), in that competition dominates in an absolute as well as in a relative sense. In contrast to Hannan and Carroll, density-dependent competition has a weaker effect on the founding rate of Italian cooperative banks in the second segment of the population than on the U.S. life insurance companies they studied.

While the two segments differ significantly in terms of density-dependent competition, they are similar in terms of legitimation. In the first segment, the founding rate at its maximum is 72 percent larger than the rate estimated at minimum density. In the second segment, the founding rate at its maximum is 78 percent larger than the rate estimated at minimum density. This result may be due simply to the effect of left-truncation in the sample (i.e., the observations only cover the latest part of the population history) or to the fact that legitimation and competition actually operate at different levels of analysis for these organizations, which are very sensitive to local conditions but are regulated and controlled at the nonlocal level by central monetary and credit authorities.

A related result concerns resource partitioning, which was found to affect entry significantly. Since previous studies found evidence of resource partitioning at the national level (Lomi, 1995) and at different local levels (Freeman and Lomi, 1994), it is possible to conclude that the effect of resource partitioning in the Italian banking industry operates similarly across different levels of analysis. Future research will have to explore in greater detail the implications of this new result for the ecological dynamics of organizational founding.

The study suffers from data limitations related to the relatively short period covered by the sample. As Hannan and Carroll (1992) pointed out, one shortcoming of earlier studies that provided discrepant findings for the effects of population density on organizational births and death rates was that data were missing about the history of the population at the lower range of density (Tucker et al., 1988; Staber, 1988; Delacroix, Swaminathan, and Solt, 1989). It is unlikely that this problem generated spurious results in this study, however, because the baseline model estimated on the pooled cross-sectional time series sample strongly supports the theory of density dependence and because the point of maximum founding rate implied by the estimates is always well within the observed range of density. Like other results produced by recent studies of organizational founding (Baum and Oliver, 1992; Baum and Singh, 1994), the present results cannot be explained in terms of left-truncation only. The effects of density maintain their predicted sign but become (statistically) weaker only when unobserved heterogeneity due to location dependence is controlled for.

In spite of these data limitations, the results of the study have broad theoretical, methodological, and substantive implications. From the theoretical point of view, the study contributes to a growing literature showing that organizational populations are internally differentiated and that vital rates vary systematically across heterogeneous segments of the population. Different researchers have used a variety of criteria for disaggregation, such as legal form and ownership structure (Freeman, 1990; Ranger-Moore, Banaszak-Holl and Hannan, 1991; Rao and Neilsen, 1992), technology (Barnett and Carroll, 1987), niche width and niche overlap (Hannan and Freeman, 1983; Baum and Singh, 1994; Lomi, 1995), evolutionary strategy (Brittain and Wholey, 1988; Tucker, Singh, and Meinhard, 1990), size (Haveman, 1993), and geographical location (Barnett and Carroll, 1987; Carroll and Wade, 1991; Swaminathan and

Wiedenmayer, 1991; Baum and Mezias, 1992; Hannan and Carroll, 1992). In this paper the organizational population was not subdivided on the basis of abstract a priori categories, but the existence and implication of heterogeneous segments were estimated directly from the data. Models estimated on the pooled cross-sectional time series sample without accounting for unobserved heterogeneity produced aggregate results that were consistent with the theory of density-dependent founding rates.

From the methodological point of view, future research should acknowledge that differences among populations and their environments affect the details of the founding process. Founding processes are sensitive to distributional assumptions, and a single parametric model of founding rates has yet to be found that is statistically best in all populations (Hannan and Carroll, 1992: chap. 4). With many possible alternatives and no compelling theoretical reason for preferring one mixing distribution over another, it is essential to minimize the impact of arbitrary distributional assumptions on inferential results. The estimation strategy adopted here ensured that the inferential conclusions reported do not hinge on arbitrary distributional assumptions about the heterogeneity components and therefore produced more reliable results.

From the substantive point of view, our understanding of the evolution of cooperative banks in Italy is enhanced by bringing the role of spatial factors back into the study of organizational dynamics. Ecological factors related to location will be of crucial importance for the survival of these specialist banks because large-scale institutional change is rapidly eroding the historical, legal, and cultural boundaries around populations of banking organizations in Europe. Whether these institutional blending processes will result in niche desegregation, how this will affect the survival chances of specialist and generalist organizations, and what agency arrangements will enjoy an evolutionary advantage under the new regime are all future research questions that have the potential to provide us with a greater understanding of the dynamics of organizational populations and communities.

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