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LOVE AND LIFE BETWEEN THE CENSUSES: A MODEL OF FAMILY DECISION MAKING IN PUERTO RICO, 1950-1960

Marc Nerlove and T. Paul Schultz

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"Varium et mutabile semper Femina."

Virgil

"I report, as man may of God's work
— all's love, yet all's law."

Robert Browning

PREFACE

The research on which this study is based was supported in pa by the Agency for International Development, Department of State, Contract csd-2151 with The Rand Corporation, and in part by the National Science Foundation under grants FSF-GS818 to the Cowles Foundation for Research in Economics at Yale University and NFS-GS2670 to the University of Chicago.

This study of the family formation process in Puerto Rico is the first of four "country studies" undertaken at Rand under AID sponsorship. The others (to follow) consider particular aspects of human fertility patterns in East Pakistan, Taiwan, and the Philippines. This study is perhaps the most ambitious of the four, given the more comprehensive and reliable data available for Puerto Rico. It is also the most tentative in its empirical results and is offered as a progress report on the feasibility and potential usefulness of comprehensive analysis of the system of family behavior that surrounds reproductive decisions. Estimates for the simultaneous-equations model developed here should ultimately provide a basis for a country-specific simulation model of short-run demographic and economic change. The purpose of this general model of family decision making is to contrast and evaluate the dynamic repercussions of alternative policy choices for the development of Puerto Rico. Much work remains to transform the analytic approach outlined in this study into a reliable guide for policy making, but the value of even this preliminary study is to illuminate the interactions among social and economic choices made by the family.

In this research, the authors benefited from the assistance in data processing and programming of John Enns and Kenneth Maurer of The Rand Corporation, and of Dennis De Tray, Alan Freiden, and Paul Weiss of the University of Chicago. The authors are indebted to Gary S. Becker, Dennis De Tray, Alan Freiden, Joseph P. Newhouse, T. W. Schultz, and Karl E. Taeuber for helpful comments and advice, not all of which have been incorporated in the present Memorandum.

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SUMMARY

Demographic phenomena such as marriage, birth, migration, and labor force participation are conditioned by economic factors and in turn contribute to the subsequent demographic and economic characteristics of the population. If, as seems likely, these types of behavior strongly interact and have roots in the common attributes of the individual and his environment, quantitative analysis should seek to specify and estimate the relevant behavioral relations simultaneously. To view one behavioral relationship at a time is likely to produce misleading estimates of cause and effect on which policy should be based and obscure the direct and indirect ways in which policy instruments influence the evolution of behavioral patterns in the household sector. This Memorandum develops a simultaneous-equations model that accounts for va demographic-economic behavior linked to the family formation proces The lack of firm empirical or theoretical guidance on the specifica of many of the behavioral relationships implies that the specific f of the model is tentative and exploratory. But, it is argued, thes types of household behavior must be analyzed within a common framew and this study demonstrates the feasibility of this approach.

The basic data for the model refer to a large number of small administrative regions in Puerto Rico, called municipios, in the ce years 1950 and 1960, as well as intervening years for some variable The model is an elaboration and extension of the single-equation an sis of birth rate determination by Schultz. We now regard fertilit as jointly determined with women's participation in the labor force, interregional migration, personal income, and the prevalence and character of marital unions. Death rates, industrial structure of employment, educational levels, and unemployment rates are treated a exogenous, but the age and sex composition of an area over time is determined by the model given the initial state.

A salient empirical finding is the strong positive relationship between birth rates and prior crude death rates lagged three and fou years. During this period the decline in mortality is contributing

substantially to the decline in fertility. The relative earnings capacity of men and women in the labor force accounts for much of the variation in marital arrangements across the Commonwealth and thereby indirectly affects fertility, labor force participation, and possibly migration. The level of personal real income is weakly but positively associated with the birth rate, holding constant for marital composition, recent mortality, adult education, and women's participation in the labor force. As anticipated, income differentials appear to be the major determinant of interregional migration flows, though the propensity to migrate differs greatly, of course, among the age and sex cohorts.

Though the specific model estimated in this study has a number of shortcomings, particularly with regard to accounting for interregional migration, the analytical approach to the household sector represents a promising instrument for extrapolating or simulating behavioral patterns for policy making and planning, with respect to both demographic and economic goals of the society.

CONTENTS

PREFACE v

SUMMARY vii

LIST OF TABLES xi

ection

I. INTRODUCTION 1
Labor Force Participation of Women 3
Family Composition -- Fertility 5
Marriage 6
Personal Income 7
Migration 9
Fertility and a Household Model of Choice 11

II. THE MODEL 13
Birth Rate Equation 24
Female Labor Force Participation Equation 27
Income Equation 29
Two Equations for Marriage 30
Migration Equations 31

III. DATA AND TECHNIQUES 33
Basic Data Sources 33
Estimation of Dynamic Relations in a Cross-Section of Time Series 37

IV. EMPIRICAL RESULTS 43
Introduction 43
Birth Rates 43
Female Labor Force Participation Rate 49
Personal Income 52
Legal and Consensual Marriages 55
Migration Equations 58

V. IMPLICATIONS FOR POLICY AND FURTHER RESEARCH 63
Feasibility and Necessity of a Simultaneous Equations Model 63
Closing and Simulating the Model 64
Some Qualitative Implications for Policy 67
Improving Individual Equations 69
Conclusion 71

Appendix
A. MIGRATION 73
B. REGRESSION RESULTS FOR 1950 AND 1960 SEPARATELY, EQUATIONS 2 THROUGH 5 83

C. REGRESSION RESULTS FOR 1950 AND 1960 COMBINED EXCLUDING SAN JUAN, EQUATIONS 2 THROUGH 5. 91
D. REGRESSION RESULTS FOR EQUATION 1 WITH INFANT DEATH RATE AS THE MEASURE OF MORTALITY REGIME. 95
E. MEANS AND STANDARD DEVIATIONS OF DERIVED VARIABLES USED IN REGRESSION ANALYSIS 97

REFERENCES. 99

SUMMARY OF DERIVED VARIABLES. 105

LIST OF TABLES

Text

1. Basic Variables and Parameters for a Demographic-Economic Model of Family Decision Making in Puerto Rico. 15
2. Derived Variables Used in the Model 18
3. List of Variables Included in the Estimated Equations 23
4. List of Puerto Rican Municipalities 35
5. Equation 1: y_1, Crude Birth Rate, 1950-1960. 44
6. Equation 2: y_2, Female Labor Force Participation Rate, 1950 and 1960 Data Combined 50
7. Equation 3: y_3, Income Per Person of Labor Force Age, 1950 and 1960 Data Combined 53
8. Equation 4: y_4, Percent of Women (14+) Currently Legally Married, 1950 and 1960 Data Combined. 56
9. Equation 5: y_5, Percent of Women (14+) Currently Consensually Married, 1950 and 1960 Data Combined. 57
10. Equation 6: $y_{13}(k)$, Female Migration in Ten Age Groups, Between 1950 and 1960 59
11. Equation 6: $y_{14}(k)$, Male Migration in Ten Age Groups, Between 1950 and 1960 60

Appendixes

A-1. Annual Net In-migration Rates, Means and Standard Deviations for 75 Municipalities of Puerto Rico and Net Migration Rates to the Mainland United States 80
A-2. Survival Rates for Puerto Rico, 1950-1960 81
B-1. Equation 2: y_2, Female Labor Force Participation Rate, 1950 Data 83
B-2. Equation 2: y_2, Female Labor Force Participation Rate, 1960 Data 84
B-3. Equation 3: y_3, Income per Person of Labor Force Age, 1950 Data 85
B-4. Equation 3: y_3, Income per Person of Labor Force Age, 1960 Data 86
B-5. Equation 4: y_4, Percent of Women (14+) Currently Legally Married, 1950 Data. 87
B-6. Equation 4: y_4, Percent of Women (14+) Currently Legally Married, 1960 Data. 88
B-7. Equation 5: y_5, Percent of Women (14+) Currently Consensually Married, 1950 Data. 89
B-8. Equation 5: y_5, Percent of Women (14+) Currently Consensually Married, 1960 Data. 90
C-1. Equation 2: y_2, Female Labor Force Participation Rate, 1950 and 1960 Data Combined, San Juan Excluded. 91
C-2. Equation 3: y_3, Income per Person of Labor Force Age, 1950 and 1960 Data Combined, San Juan Excluded 92
C-3. Equation 4: y_4, Percent of Women (14+) Currently Legally Married, 1950 and 1960 Data Combined, San Juan Excluded . . 93
C-4. Equation 5: y_5, Percent of Women (14+) Currently Consensually Married, 1950 and 1960 Data Combined, San Juan Excluded . . 94

D-1. Equation 1: y_1, Crude Birth Rates with Infant Death Rates, z_9 as Measure of Mortality Regime 95
E-1. Means and Standard Deviations of Derived Variables Used in Regression Analysis 97

I. INTRODUCTION

The family is commonly encountered as the unit of observation for the study of savings and expenditure behavior. Recent reinterpretations of the theory of consumer choice in terms of the allocation of "time and goods over time" have further emphasized the interaction between the individual and the family over the life cycle (Becker [1965, 1967], Ghez [1968]). In addition to being a consumer unit, the family is a supplier of labor. It affords the traditional means for greater individual specialization among market, nonmarket, reproductive, and child rearing activities. Generational cycles in both physical savings and human capital formation in children and adults further constrain and shape consumption opportunities of persons within or potentially within a family unit (Ben-Porath [1967], Ando and Modigliani [1963], Tobin [1967]). Some of these aspects of behav have been investigated, but generally each separately, neglecting their jointly determined character. Demographic phenomena such as marriages, births, migration, and labor force participation are conditioned by economic factors and in turn contribute to the subsequent demographic character and economic characteristics of the population. If, as seems likely, these types of behavior strongly interact and have roots in common attributes of the individual and his environment, quantitative analysis should seek to specify and estimate the relevant behavioral relations simultaneously.

This Memorandum develops a simultaneous-equations model that accounts for various demographic-economic behavior linked to the family formation process.[1] The basic data for our model refer to a large number of small geographical regions in Puerto Rico, called

---

[1] Frequently each of two behavioral decisions will depend upon tl other. Generally each contains a random element -- that is, the dec: sions are not deterministic. Furthermore, these random elements wil: usually not be independent of one another. To avoid biased estimate: of the underlying cause and effect relationships in this case, speci: simultaneous-equations estimation techniques are required (see J. Jo [1963, Chap. 9]). These statistical methods have been used in econo for several decades, and are found increasingly today in quantitativ sociology and psychometrics (Goldberger [1969, 1970]).

municipios, in the census years 1950 and 1960 as well as intervening years for some variables. Our model is an elaboration and extensio of the single-equation analysis of birth rate determination by Schu [1969a]. We now regard fertility as jointly determined with women' participation in the labor force, interregional migration, personal income, and the prevalence and character of marital unions. Death rates, industrial structure of employment, educational levels, and unemployment rates are treated as exogenous, but the age and sex co position of the population of an area over time is determined by th model given the initial state.

In this introductory section we review several literatures per taining to behavior we seek to model. In Section II we discuss the specification of our model. In Section III we report some prelimin statistical results from estimating the model for contemporary Puer Rico. In Section IV we summarize and interpret our findings and describe a program for further research designed to improve and com plete the model.

To predict a particular class of behavior of women is essentia the task of this study; that is, we ask what factors affect women's choice of marital status (single, consensually married, legally mar labor force status (in or outside of the civilian labor force) and eproductive status (currently bearing a child or not). Although omplicity with men may be assumed in this behavioral scheme, it is he observable behavior of women on which we focus; men, regardless f marital status and family composition, are commonly found in the abor force, earning what they can.[1] To a great extent, the level-

---

[1]Many of the relationships we ultimately estimate can be view artially reduced forms for both the appropriate supply and demand chedules. For example, in the marital status equations female de nd male supply of marital agreements (or vice versa) require impl: arket prices and quantities. We refer to a locus of equilibria t ssumed to be a function of the costs and returns to marriage in e: f two forms for both of the sexes. In the female labor force par ation equation, our inability to identify and estimate both the f labor supply and derived demand schedules leads us to the traditio egression equation that includes determinants of shifts in both t lemand and supply schedules.

market income, the amount and composition of home produced and consumed goods and services, and the family composition are contingent upon the woman's activities. We assume here that the extent to which children are set to work or sent to school is exogenous to the model, limited only by the availability of schooling in the region. Within this behavioral complex the woman's decision whether to participate in the labor force has perhaps received the most systematic attention, particu larly in the United States.

LABOR FORCE PARTICIPATION OF WOMEN

Analyses of participation rates of women often focus separately on single and married women and implicitly assume the prevalence and duration of marriage influence the size of the female labor force. It is also presumed that extending the years of adolescent schooling exerts a direct effect on participation by delaying the age when a woman may conveniently enter the full-time labor force (Bowen and Finegan [1965]). The effects of labor market conditions and household opportunities for economic and reproductive activity require a more detailed description.

Women are more lil
employment opportunities are attractive: that is, when the rewards are large and certain, alternative economic resources are scarce, and family responsibilities are compatible with employment. These general expectations are confirmed by cross-sectional studies of female labor force participation in the United States, although the form of the statistical model and the measurement of some of the variables cause certain difficulties. The earnings of women are directly associated and unemployment is inversely associated with female participation rates (Mincer [1960, 1962, 1966], Cain [1966, 1967], Bowen and Finegan [1965, 1969]).[1]

[1]Although cross-sectional estimates may overstate the magnitude of unemployment's discouraging effect on labor force participation rates of secondary workers, the direction of the net effect in U.S. studies appears unambiguous (Mincer [1966]). To our knowledge similar studies have not been undertaken in low-income developing countries to estimate the relationship between female labor force participation

Education may also plausibly affect women's participation by increasing her employability (opportunity income) and perhaps also by increasing her desire to work. Despite collinearity between female earnings and education both variables appear to exert independent effects on female participation according to cross-sectional studies of the United States (Bowen and Finegan [1965]). But the effect of education on participation is not invariant. In the United States, holding constant for age, color, and income, differences in participation rates for white wives associated with their education attainment nearly vanish if one of their children is less than six (Sweet [1968]). In this case, we might conclude that educated wives are strongly influenced in their participation behavior by the demands of rearing their own young children. The evidence from other countries on the effects of education on participation is more limited.[1]

The scarcity or adequacy of other resources available to the household may be measured by family income minus that amount the woman earns herself. This measure of "other income" is inversely associated with U.S. participation, and more refined measures of income adequately adjusted for the consumption requirements of the family's composition strengthen this inverse association (Sweet [1968]).

Regional differences in industrial structure may affect the derived demand for the labor services of women and hence influence employment opportunities, wages, and participation. An index of sectoral demand

rates and earnings opportunities, taking account of the interdependence of the participation decision with decisions related to marriage and reproduction.

From the 1960 Census of Puerto Rico in which personal income by sex is tabulated by municipality, we found a significant positive association between female participation rates and female income, and negative association between female participation and both male income and unemployment. The lack of similar sex-specific income data by region from the 1950 Census precludes our investigation of this relationship in later empirical analyses.

[1] Harman (1969) in an analysis of individual survey data for the Philippines finds female labor force participation strongly associated with age, directly with educational attainment, and inversely with marriage.

for female labor services is associated cross-sectionally in the United States with married female participation rates (Belloc [1950], Bowen and Finegan [1965]).[1]

FAMILY COMPOSITION -- FERTILITY

The composition of a woman's family is strongly associated with her participation in the U.S. labor force. Typically, the number of own children under the age of 18 and age of youngest child are both strong predictors of participation, and have been interpreted by Cain, for example, as measures of the opportunity value of a mother's time in the home (Cain [1965]). But can this variable be treated as exogenous? (Sweet [1968].) Because birth control is widely practiced in the United States (and in Puerto Rico),[2] family composition reflects in part past decisions whether to participate in the labor force today or bear and rear children today; hence, current family composition and participation are largely simultaneously and jointly determined in an earlier planning period, though each is subject to some uncertainty for the individual household.

The relationship between the recent past fertility and current participation of women in the labor force can also be interpreted as partly biological in origin. To some extent sterile and subfecund women may be found in the labor force because they cannot bear children. On the other hand, the desire to work may lead to family restriction, and conversely, the desire for a large family may lead to restricted labor force participation. Probably both the biological and behavioral mechanisms operate in all societies, but in low-income countries where family planning is less reliable and widespread we might give the biological explanation somewhat greater weight (Stycos [1968], Weller [1968, 1970]). Nonetheless, even in the United States,

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[1] Teenage female participation rates appear less responsive in the postwar period to this type of demand index, but this may only reflect the growing segmentation of the U.S. labor market, in which teenage employment opportunities are in different sectors than adult. Our own results indicate a positive cross-sectional association between such an index and the overall rate of female labor-force participation.

[2] Presser [1969].

few studies have proven that women's career ambitions and their desire to work or to undertake nonfamily activities actually contribute to lower desired or achieved fertility (Pratt and Whelpton [1958], Ridley [1959], Westoff [1961, 1963], Gendell [1967], Sweet [1968]). It may be more accurate to suppose that participation primarily affects the timing of family formation rather than the completed size (R. Freedman and Coombs [1966a, 1966b], D. Freedman [1963]). Causality is hard to confirm among the interactions, for better educated women can earn more, do earn more, marry later, and have their fewer children closer together, permitting these women to stay in the labor force for more years.

With reference to low-income countries the causal inferences relating female labor force participation and fertility are further complicated by inadequate measures of labor force participation. The importance of cottage industries, home handicraft, and live-in domestic service employment for women in less developed countries suggest that all classes of female employment may not be equally incompatible with concurrent child bearing and rearing. Only occupational commitments of a more modern sort, outside of the home, constrain the pursuit of traditional family roles and are therefore less often observed in conjunction with high fertility.[1]

MARRIAGE

Once out of school the young woman may have a variety of options including marriage, labor force participation, and perhaps migration. What factors influence her choice, or the choice of the household of which she is a member? Several studies have shown that over time in several developed countries there exists an association between the business cycle and the timing of marriage and subsequent birth rates

[1] When this distinction among types of female employment is made, fertility is generally found to be lowest for women in modern occupatic (those requiring work away from home, for example, clerical and technical activities), intermediate for domestic services and handicrafts, an highest for women not classed as participants in the labor force (Hatt [1952], Jaffe [1959], Jaffe and Azumi [1960], Gendell [1967], Weller [1968, 1970], Collver and Langlois [1962]). Exceptions nevertheless occur in some cultural settings (Stycos [1968], Gendell and Kantner [1969]).

(Silver [1965, 1966]). Some evidence exists of cyclical fluctuation in marriage and birth rates in low-income countries, as for example Puerto Rico (Combs and Davis [1951], Hatt [1952], Myers and Gibson [1968]), but cross-sectional differences in the prevalence of marria stem largely from the secular rise in the age of marriage. These cross-sectional differences may be directly responsive to the level of personal income and inversely to unemployment, but more modestly than marriage rates and birth rates, for they represent differences in a stock rather than differences in the adjusting flow over time.

Where different forms of marriage are common, as in Puerto Rico where consensual and legal unions are distinguished, the choice of marital status becomes more complex: What factors lead the partners to prefer marriage and what form of marriage serves the mutual and the separate interests of the partners?[1] The alleged instability of consensual unions is frequently thought to be more of a disadvantage to the woman than to the man. But in the urban lower-class environment where this institution is common, observers have noted that consensual unions are often matriarchal; the woman, because of her positi in the labor force and the man's frequent exposure to unemployment, appears to be both the economic and social authority opting for the consensual union (Lewis [1965], Goode [1963]). If this hypothesis is correct, conditions in the labor market that strengthen the woman's economic status relative to the man's should be associated with more frequent adoption of the consensual marriage route, at least during the initial years of marriage until the man has established himself economically. This hypothesis remains, however, a subject of controversy in the sociological literature (Goode [1963]).

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[1]"About a quarter of all unions in Puerto Rico are consensual, and the persistence of the pattern indicates its utility. Although marriage is generally viewed as more desirable than consensual union, the latter is not seriously condemned by the lower-class community and has a sufficient number of advantages to make it attractive. Because of the insulation of the sexes before marriage, men and women enter this status in considerable ignorance of one another. Divorce for this class is difficult and costly. The consensual union serves as a kind of trial marriage in which a mistake can be corrected without undue hardship." Stycos [1955, p. 245].

PERSONAL INCOME

The level of potential income in a community is a function of its supply of economic resources and its demands for these resources. Income from labor constitutes generally from half to three-fourths of national income, and, since labor income is reported more completely in surveys and censuses, the larger part of measured income is likely to be attributed to labor. Formal scnooling is one form of human capital that is readily measured and presumed to enhance the productivity of labor and hence to raise its earnings potential. But the earnings potential for comparable labor services may not be uniform across regions or sectors, for the costs of geographic and sectoral mobility and imperfections of information hinder the movement of labor to eliminate labor market disequilibria.

Unemployment is the most obvious indication of an insufficient level or mix of local derived demand for labor, given the existing supply and institutional structure. Widespread unemployment should therefore depress potential returns to labor in general. Unemployment means fewer persons in the labor force are earning an income; and, as indicated above, unemployment probably also dissuades some from entering the labor force altogether, adding further to the inverse association expected between unemployment and personal income per adult person.

Sharp sectoral shifts in the composition of derived demand for labor, sometimes associated with the development process, are another source of differences in the returns to otherwise comparable labor. The shift of workers out of agriculture is an important feature of development associated with rising incomes.[1] Such a rate of exodus from agriculture suggests substantial sectoral disequilibrium and lower potential earnings of labor in agriculture in comparison with

[1] In Puerto Rico the proportion of the labor force in agriculture fell from 39 percent in 1950 to 27 percent in 1960. Unfortunately the census definition of the farming population changed between the censuses of 1950 and 1960 adding further to possible errors in the measurement of this variable.

other sectors of the economy. Other studies of personal income determination in developing countries confirm the effect of education and the agricultural disequilibrium on personal incomes. But unemployment, on the other hand, perhaps because of its vague nature in the low-income setting, does not appear always to be strongly associated with depressed personal incomes (Harman [1969]).

The observed level of personal income per adult in a community is not only a function of the determinants of personal earnings potential but is also a function of how persons allocate their time, or, more central to our model, how women allocate their time between labor force participation and homemaking and child rearing activities. Childbearing should be negatively associated with the level of persona incomes, and, conversely, greater participation of women in the labor force, other things equal, should be associated with greater personal income per adult in a community. Similarly, the greater the extent to which children add their potential earnings to those of the family, at a cost to their schooling or leisure, the greater should personal incor per adult be. Since labor force participation of both women and children tend to be inversely associated with income or income of the head of household, it may be difficult to observe directly the net contribution of these two groups of secondary workers to personal income (Bowe and Finegan [1965], Cain [1966], Sweet [1968]). It is possible that the prevalence of marriage and the frequency of childbirth are better predictors of the opportunity costs of withdrawing women from the labo force than their actual participation rates.

Home production activities of families have been studied from the point of view of having children. Household behavior in this sphere is not investigated here for lack of adequately disaggregated income data.

MIGRATION

Viewed as an activity with associated costs and returns to the productivity and earnings of the human agent, migration has been analyzed as one form of human capital formation (Sjaastad [1962]). The net cost and benefit streams, broadly defined, may be reduced to

a single rate of return, which is presumed to be associated with the actual migration rate. This view of migration as an investment activity undertaken when discounted benefits outweigh costs is conceptually attractive, but it requires data for empirical verification that are not commonly available, that is, future earnings profiles in various locations for various persons and prospective direct and opportunity costs associated with migration. For this reason, we suspect, the rate of return approach to migration has received scant direct verification and has been "tested" only in terms of associations between migration and linear combinations of factors thought likely to influence the rate of return to, or attractiveness of, migration (Diehl [1966]). An alternative literature disjointly deals with the determinants of migration as either "push or pull" factors and does not explicitly focus on the differences between opportunities in a locality and elsewhere versus the costs of migration as the underlying determinants of migration (Lee [1966]). Both approaches may be relevant to understanding migration.

It is not now empirically feasible to calculate the "expected" rate of return to migration as a predictor of the migration rate. Analysis of migration may nevertheless proceed as follows: Aggregate variables that should affect the rate of return to migration, such as regional differences in earnings, unemployment, and the structure of employment, are first identified. These may in turn be related to migration for different age-sex categories in order to distinguish how the composition of the population influences the responsiveness of individuals to such differences among regions.

It is well established that particular groups in a population may be more susceptible than others to migration opportunities: the young, who are still relatively unencumbered and flexible and look forward to many rewarding years in a new location; the better educated, who are better equipped to evaluate the opportunities and uncertainties abroad; and often the women, who if equally educated with men may have more to gain and less to lose from leaving traditional rural society and accepting the responsibilities and opportunities of progressive urban society. To the extent that the composition of regional

populations differ with regard to age, sex balance, and educational attainment, one might expect them to differ in their susceptibility to migration given the same differential opportunities. There are two methods for dealing with these compositional qualities of the population. Either the compositional variable may be added to the predictive model, implicitly assuming a particular functional relationship between the characteristic and the probability of migration, or migration may be considered separately within specific age-sex-education classes without specifying the form of the relationship between the compositional characteristics and the probability of migration. Since the age and sex composition of the population is influenced by past migration and influences other aggregate behavioral patterns we wish to account for, a disaggregated approach to migration is necessary here to capture the dynamic properties of our system of behavioral relations.

FERTILITY AND A HOUSEHOLD MODEL OF CHOICE

Reproductive behavior is not always viewed as responsive to man's environment. The Malthusian premise -- that man procreates at a relatively constant rate and mortality changes to achieve reasonable balance between population and environment -- permeates many investigations of fertility and population growth. When historical trends in population and economic growth coincide, scholars cannot agree whether this is coincidental, or if not, which is cause and which effect (Habakkuk [1953], Ohlin [1955], Clark [1967]). In contemporary high-income countries there is abundant evidence that an association exists between the business cycle and birth rates (Silver [1965, 1966]), but evidence of long-run swings in economic activity linked to cycles in population growth and fertility is more controversial (Abramovitz [1964]). A theory for these longer period fluctuations in U.S. fertility has been recently advanced by Easterlin [1968], but as yet econometric techniques have not been used to test this approac with disaggregated data

Aggregate econometric models have tended to neglect the economic and demographic behavior of the household sector. The only major

effort to study micro behavior of the household is that of Orcutt and his associates [1961]. Orcutt et al., however, do not attempt to estimate simultaneously behavioral relations characterizing household decisions; rather they seek to estimate the probability of specific events -- such as marriage, birth, female labor force participation -- given a set of conditioning features of the population or "status variables," typically demographic and economic in nature. Realistic simulation of demographic and economic change at the household level is their goal, as it is ours, but our methodologies differ.

Becker [1960] first proposed to analyze fertility as the outcome of consumer choice, but his efforts to test his theory empirically were limited. Examination of the empirical implications of this approach, generalized to conform to the setting of the low-income country, has since yielded stronger confirmation of the model. Abstracting from uncertainty implicit in the production of children, because of both biological factors and the unreliability of birth control, earlier work has accounted for a significant part of the variation in fertility interregionally and over time in terms of the environmental constraints on parent demand for children. Of substantial importance in past work is the compensating association between the frequency of (child) death and birth in families (Harman [1969]) and in communities (Schultz [1969a, 1969b, 1970]). The object of the next section is to outline a more comprehensive model of household behavior in which fertility is treated as jointly determined with the other decisions and their consequences.

THE MODEL

In this section, we formulate a model of several interdependent household decisions that explicitly takes into account the simultaneity involved in these areas of choice. The key variables in our study of Puerto Rican demographic and economic structure in the 1950s are (1) family size, (2) female labor force participation rate (3) marriage patterns, (4) personal income, and (5) migration from rural to urban areas. Desired or even realized family size is not, unfortunately, an observed variable. Consequently, we have chosen to focus on a directly observable variable, the crude birth rate.

Whether or not a woman chooses to work and at what points durin her life she does so are related to the size and composition of her family over time. The presence of a number of small children in the household may effectively preclude participation in the labor force away from home. Fertility and labor force participation are also affected by marriage, and marriage decisions are made jointly with decisions affecting ultimate family size and timing as well as labor force participation. In Puerto Rico the distinction between legal marriage and consensual union is a demographically important one, since fertility and labor force participation rates differ between the two types of households. Proportions of women in the relevant age groups who are legally married or in consensual unions are thus also endogenous to the model.

Migration from rural to urban areas within Puerto Rico and betwe Puerto Rico and mainland United States is highly age and sex specific In our model we assume migration depends upon economic differences, such as income and employment opportunities, that exist among regions and on the educational attainments of potential migrants. Such migra tion certainly affects the crude birth rate and possibly other demographic variables through its differential impact upon the age-sex composition in areas from which and to which migration occurs. Migra tion may also selectively change the behavior composition of populati even within a given age-sex cohort. Demographic variables, in turn,

affect migratory flows by affecting the composition of the populations from which these flows take place.

In addition to the crude birth rate, the female labor force participation rate, and migration flows, the endogenous variables of our model explicitly include the proportions of women legally and consensually married and a variable reflecting income both in cash and in kind, as well as, implicitly, the total population and its age-sex composition. In the present version of the model, we do not explicitly state the identities relating the population in various age-sex categories to migration rates, birth rates, and death rates necessary to close the model demographically. In simulations with the model such closure is necessary, and the appropriate identities will be explicitly incorporated. Since age-specific death rates are not available by the small geographical areas that are our units of observation, some of the equations needed to close the model demographically will not be identities in the strict sense of the word but only empirical approximations thereto.

Our birth rate equation is estimated using data for 75 municipios over the eleven-year period 1950-1960; the migration equations are fitted to purely cross-sectional data on the 75 municipios; and the remaining equations are estimated by combining the data on the 75 municipios for the two census years 1950 and 1960.

Mnemonic symbols and descriptions of the main endogenous and exogenous variables in our analysis are given in Table 1. The number of endogenous variables exceeds the number of equations explicitly formulated because of the inclusion of several demographic variables only implicitly determined. The derived variables actually used in the statistical analysis are presented in Table 2. The equations estimated are summarized in Table 3. On the last page of this Memorandum is a fold-out summary table of the derived variables in the complete model for the reader to consult as needed.

Estimates of the parameters of the equations listed in Table 3 are presented in the next section. Here we present some justification

Table 1

BASIC VARIABLES AND PARAMETERS FOR A DEMOGRAPHIC-ECONOMIC MODEL OF FAMILY DECISION MAKING IN PUERTO RICO

Symbol	Description
b_{it}	Total births to mothers residing in municipio i in period t.
fl_{it}	Number of women 14 years or older economically active in municipio i in period t.
I_{it}	Personal income in cash and kind in 1959 dollars in municipio i in period t.
mc_{it}	Number of women living in free or in consensual unions in municipio i in period t.
ml_{it}	Number of women legally married in municipio i in period t.
$Mf_{it}($	Net in-migration (negative for out-migration) of females in age group k to municipio i in period t.
$Mm_{it}($	Net in-migration (negative for out-migration) of males in age group k to municipio i in period
$Nf_{it}($	Female population in age group k in munic period t.
$Nm_{it}($	Male population in age group k in municipio i in period t.
N_{it}	Total population in municipio i in period t.
$c_{it}(k$	Number of children in age group k in school in municipio i in period t.
d_{it}	Total number of deaths of persons residing in municip i in period t.
IM_{it}	Number of deaths of infants less than one year of age residing in municipio i in period t.
DUM	Dummy variable: 0 in 1950, 1 in 196(
$E_{it}($	Total employment in industry j in municipio period t.[a]

Table 1 (continued)

Symbol	Description
fw_{it}	Number of unpaid family workers in municipio i in period t.
LF_{it}	Total civilian labor force in municipio i in period t.
LFA_i	Civilian labor force in agriculture in municipio i in period t.
s_{it}	Median years of schooling for persons 25 years or over in municipio i and period t.
U_{it}	Total unemployment in municipio i in period t.
α_k	Age-specific birth rate per thousand women in age grou k for Puerto Rico as a whole in 1954.[b]
β_j	Proportion of total employment in industry j that is female for Puerto Rico as a whole, mean ratio for 1950 and 1960.[a]
γ_k	Age-specific female labor force participation rates in age group k for Puerto Rico as a whole: mean rate for 1950-1960.[c]
δ_k	Age-specific legal marriage rates for age group k in Puerto Rico as a whole; mean rate for 1950-1960.[c]
ε_k	Age-specific consensual union rates in age group k for Puerto Rico as a whole; mean rate for 1950-1960.[c]

Notes:

ndustry groups and the corresponding βs are:

		β_j
1.	Agriculture	018
2.	Forestry and fishing	015
3.	Mining	012
4.	Construction	015
5.	Nondurable manufacturing	525
6.	Durable manufacturing	226
7.	Transportation, communications, and utilities	057
8.	Wholesale trade	111
9.	Retail food and dairy products	084
10.	Eating establishments	190

Notes to Table 1 (continued)

11.	Other retail	.255
12.	Finance	.310
13.	Business and repair servic	.038
14.	Private housing services	.867
15.	Other personal services	.333
16.	Entertainment	.183
17.	Hospitals, medical and professional services	.552
18.	Educational services, public	.728
19.	Educational services, private	.710
20.	Public administration	.251
21.	Industry not reported	.417

[b]The values of α_k are:

k	ge Group	
1	0-14	0.6
2	15-19	97.5
3	20-24	254.6
4	25-29	218.9
5	30-34	159.9
6	35-39	131.3
7	40-45	47.5
8	over 45	10.8

[c]The values of γ_k, δ_k, and ε_k in percent are as follows

	γ_k	δ_k	ε_k
15-19	13.3	14.7	4.2
20-24	30.5	51.1	11.3
25-29	28.3	70.7	13.9
30-34	27.3	75.8	14.4
35-39	25.6	76.5	15.3
40-44	24.0	73.3	13.4
45-49	21.0	69.6	11.5
50+	10.6	42.7	5.8

Table 2

DERIVED VARIABLES USED IN THE MODEL
(municipio and period subscripts suppressed)

Symbol	Description and Formula
1. Endogenous Variables	
y_1	Crude birthrate (per thousand): $(b/N) \cdot 10^3$
y_2	Female labor force participation rate (percent): $[fl/Nf(14+)] \cdot 10^2$
y_3	Personal income in cash or in kind per person of labor force age (in hundreds of 1959 dollars): $I/[Nf(15-65) + Nm(15-65)]$
y_4	Percentage of females 14 years and over legally married (percent): $[ml/Nf(14+)] \cdot 10^2$
y_5	Percentage of females 14 years and over living in consensual unions (percent): $[mc/Nf(14+)] \cdot 10^2$
y_6	Index of female age composition, age-specific birth rate weights (per thousand): $\sum_{k=1}^{8} \alpha_k \, Nf(k) / N$ k = 0-14, 15-19, 20-24, 25-29, 30-34, 35-39, 40-44, over 45
y_7	Sex ratio in marriageable age group: $Nm(15-34)/Nf(15-34)$
y_8	Children 0-4 per woman 15 - 49 $[Nf(0-4) + Nm(0-4)] / Nf(15-49)$

Table 2 (continued)

Symbol	Description and Formula
y_9	Index of female age composition, age-specific labor force participation rate weights: $\sum_{k=1}^{8} \gamma_k \, Nf(k) \Big/ Nf(14+)^a$
y_{10}	Adult sex ratio: $Nm(15\text{-}64) / Nf(15\text{-}64)$
y_{11}	Index of female age composition, age-specific legal-marriage rate weights: $\sum_{k=1}^{8} \delta_k \, Nf(k) \Big/ Nf(14+)^a$
y_{12}	Index of female age composition, age-specific consensual marriage rate weights: $\sum_{k=1}^{8} \varepsilon_k \, Nf(k) \Big/ Nf(14+)^a$
$y_{13}(k)$	Net in-migration (observations on individual municipios) for females in age group k per capita resident females in age group k (annual compound rate): $[Mf(k)/\overline{Nf}(k)]^b$
$y_{14}(k)$	Net in-migration (observations on individual municipios) for males in age group k per capita resident males in age group k (annua compound rate): $[Mm(k)/\overline{Nm}(k)]^b$
$y_{15}(k)$	Weighted income differences among municipio for use in equation explaining $y_k(k)$, 1959 dollars: $\sum_{i=1}^{76} \left[\frac{\overline{Nf}(k) + \overline{Nf}_i(k)}{\overline{Nf}(k)} \right] (\overline{I}_i - \overline{I})^c$

Table 2 (continued)

Symbol	Description and Formula
$Y_{16}(k)$	Weighted income differences among municipios for use in equation explaining $y_{14}(k)$, 1959 dollars: $\sum_{i=1}^{76}\left[\frac{\overline{Nm}(k) + \overline{Nm}_i(k)}{\overline{Nm}(k)}\right](\overline{I}_i - \overline{I})^c$
$y_{17}(k)$	Weighted differences in unemployment rates for use in equation explaining $y_{13}(k)$: $\sum_{i=1}^{76}\left[\frac{\overline{Nf}(k) + \overline{Nf}_i(k)}{\overline{Nf}(k)}\right](\overline{z}_{7i} - \overline{z}_7)^c$
$y_{18}(k)$	Weighted differences in unemployment rates for use in equation explaining $y_{14}(k)$: $\sum_{i=1}^{76}\left[\frac{\overline{Nm}(k) + \overline{Nm}_i(k)}{\overline{Nm}(k)}\right](\overline{z}_{7i} - \overline{z}_7)^c$
$y_{19}(k)$	Weighted differences in the index of demand for female labor service for use in equation explaining $y_{13}(k)$: $\sum_{i=1}^{76}\left[\frac{\overline{Nf}(k) + \overline{Nf}_i(k)}{\overline{Nf}(k)}\right](\overline{z}_{3i} - \overline{z}_3)^c$
$y_{20}(k)$	Weighted differences in the index of demand for female labor services for use in equation explaining $y_{14}(k)$: $\sum_{i=1}^{76}\left[\frac{\overline{Nm}(k) + \overline{Nm}_i(k)}{\overline{Nm}(k)}\right](\overline{z}_{3i} - \overline{z}_3)^c$
$y_{21}(k)$	Relative population index for use in equation explaining $y_{13}(k)$: $\sum_{i=1}^{76}\frac{\overline{Nf}(k) - \overline{Nf}_i(k}{\overline{Nf}(k)}$

Table 2 (continued)

Symbol	Description and Formula
$y_{22}(k$	Relative population index for "e equation explaining $y_{14}(k)$: $\sum_{i=1}^{76} \frac{\overline{Nm}(k) - \overline{Nm}_i(k)}{\overline{Nm}(k)} c$
$y_{23}(k$	Dummy variable for use in equation explaining $y_{13}(k)$ to account for extra cos of moving to U.S.: $\frac{\overline{Nm}(k) - \overline{Nm}_{76}(k)}{\overline{Nm}(k)} c$
$y_{24}(k$	Dummy variable for use in equation explaining $y_{14}(k)$ to account for extra cost and unspecified benefits of moving to U.S.: $\frac{\overline{Nf}(k) - \overline{Nf}_{76}(k)}{\overline{Nf}(k)} c$

2. Exogenous Variables

z_1	Percentage of children aged 7-13 in school (percent): $c(7\text{-}13)/[Nf(7\text{-}13) + Nm(7\text{-}13)] \cdot 10^2$
z_2	Percentage of children aged 14-17 in school (percent): $c(14\text{-}17)/[Nf(14\text{-}17) + Nm(14\text{-}17)] \cdot 10^2$
z_3	Index of demand for female labor services: $\left\{ \frac{\sum_{j=1}^{21} \beta_j E(j)}{\sum_{j=1}^{21} E(j)} \Big/ \frac{\sum_{i=1}^{75} \sum_{j=1}^{21} \beta_j E_i(j)}{\sum_{i=1}^{75} \sum_{j=1}^{21} E_i(j)} \right\}$ Note that U.S. is not included as the 76th municipio in summations in the denominator of the above expression.

Table 2 (continued)

Symbol	Description and Formula
z_4	Adult schooling (years): s
z_5	Crude death rate (per thousand) $(d/N) \cdot 10^3$
z_6	Unpaid family workers as a percent o civilian labor force (percent): $(fw/LF) \cdot 10^2$
z_7	Unemployment rate (percent): $(U/LF) \cdot 10^2$
z_8	Percent of civilian labor force in agriculture (percent): $(LFA/LF) \cdot 10^2$
z_9	Infant death rate (per thousand live births): $(IM/b) \cdot 10$
DUM	Dummy Variable: 0 in 1950, 1 in 1960.

Notes:

[a] k = 14-19, 20-24, 25-29, 30-34, 35-39, 40-44, 45-49, and 50+.

[b] Bar denotes mean value for age group k in 1950 and age group k + 10 in 1960.
k = 0-4, 5-9, 10-14, 15-19, 20-24, 25-34, 35-44, 45-54, 55-64, 65+.

[c] Bars denote mean values for age group k in 1950 and age group k + 10 in 1960. Summation is over all municipios; unsubscripted variables refer to municipio of observation.
k = 0-4, 5-9, 10-14, 15-19, 20-24, 25-34, 35-44, 45-54, 55-64, 65+.

Table 3

LIST OF VARIABLES INCLUDED IN THE ESTIMATED EQUATIONS
(note: $x(-\tau)$ denotes x_{it}

(1). Birth rate equation

y_1: y_2, y_3, y_4, y_5, y_6; z_2, z_4, $z_5(-2)$, $z_5(-3)$, $z_5(-4)$, $z_5(-5)$, z_6

and also omitting z_5s and including z_9s.

Continuous data, 1950-1960.

(2). Female labor force participation equation

y_2:y_3, y_4, y_5, y_8, y_9; z_2, z_3, z_7, DUM

also omitting y_4 and y_5 .

(3). Income equation

y_3:y_2, y_4, y_5, y_8, y_{10}; z_4, z_7, z_8, DUM

also omitting y_2 .

(4). Legal marriages equation

y_4:y_3, y_7, y_{11}; z_2, z_3, z_7, DUM

(5). Consensual marriages equation

y_5:y_3, y_7, y_{12}; z_2, z_3, z_7, DU

(6). Migration equations (note: There are 20 equations, one for each of the age-sex categories considered).

$y_{13}(k)$: $y_{15}(k)$, $y_{17}(k)$, $y_{19}(k)$, $y_{21}(k)$, $y_{23}(k)$

$y_{14}(k)$: $y_{16}(k)$, $y_{18}(k)$, $y_{20}(k)$, $y_{22}(k)$, $y_{24}(k)$

For estimation purposes in this model we treat y_{15}, y_{16}, y_{17}, y_{18}, y_{19}, y_{20}, y_{21}, y_{22}, y_{23}, and y_{24} as if they were exogenous

for our choice of variables to be included in or excluded from the equations of the model.

BIRTH RATE EQUATION

An extensive discussion of some of the factors underlying the form of the birth rate equation used in this model may be found in Schultz [1969a]. Essentially our equation is designed to reflect a family's desire for a certain ultimate size, that is, for a specific number of surviving children as modified by the adjustments necessary as the result of deaths among offspring, actual or anticipated in the future; the intrinsic fertility of the woman as determined by her age; and certain sociological factors connected with the form and age of marriage in the community. Neither age nor form of marriage appeared to exert significant effects on the birth rate in Schultz's earlier study [1969a]; as we shall see, however, the measures used here do account for a significant part of the variation in crude birth rates interregionally and over time.

The relation between birth rates and the variables we assume to enter Equation (1) may conveniently be described in terms of the way in which some variables affect the desired ultimate family size, and then how other variables affect the way in which these desires are translated into birth rates. To the extent that children are a superior good, higher family income will, other things being equal, result in larger desired family sizes. However, it is possible that the costs of rearing a child also rise sharply with income. If costs rise sufficiently fast with income, larger incomes will be associated with smaller desired family sizes. Conversely, the more slowly costs of rearing children rise as income rises, the more likely that larger desired family sizes will be associated with high incomes. To the extent children represent provision for security in old age, or can be used in the productive process at any early age, high incomes may tend to be associated with smaller desired family sizes. The effect of income upon birth rates cannot, therefore, be predicted *a priori*.[1]

[1] It is possible, in some agricultural communities in which children play an important role as a productive input, that increasing income will

The extent and duration of schooling for children represent a readily distinguishable component of the cost of child rearing. We expect it to be inversely associated with desired family size. In general, children's schooling reduces their role in family production activities including their time to tend to younger siblings may also impose additional actual support costs on parents.

The more highly educated the parents, the higher the opportunity cost of their time; and time is a substantial component of the cost of child rearing. In addition, better educated parents may have greater aspirations for their offspring and consequently anticipate higher child rearing and schooling costs; hence, high adult education levels should be associated with small desired family sizes. Adult education may further facilitate the parents' achievement of their family size goal by improving their access to modern and more reliable methods of birth control.

Participation of women in the labor force reflects the decision of women and their families to allocate more of their time to labor force activity and less to homemaking and child rearing activities. Where the balance of incentives favors greater female labor force participation, we anticipate also that smaller family sizes will be desired.

To the extent that consensual unions are considered by those who participate in them to be more unstable than legal marriage, the typical family size, realized and desired, should tend to be smaller the more widely prevalent are free unions (Stycos [1955]). Conversely high rates of legal marriage should be associated with larger family sizes. Because the form of marriage may, in such countries as Puerto Rico, be highly age specific, however, great care must be used in the interpretation of the relation between the form of marriage and the crude birth rate. The typical pattern may be for a woman to live for

lead to increased savings in the form of investment in durable factors of production, of which children are one form.

some years, possibly her most fertile ones, in consensual union before legalizing the marriage (Hatt [1952]).

Given desired family size levels, a number of important factors affect observed birth rates. Perhaps the most important of these is the age of the woman; or, since we are dealing with a cross-section of geographical areas the characteristics of which are changing over time owing to migration, the age composition of the women in each area. Areas with a high proportion of both very young and older women will tend to exhibit low birth rates relative to areas with a high proportion of women in the most fertile age groups even though desired family sizes may be equal, since the younger women will not yet have begun their families and the older women will have completed theirs. Unfortunately, age-specific birth rates are not available by the geographical areas used in our analysis; we have, therefore, created an index of the female age composition for each municipio and year, using as weights Commonwealth-wide age-specific birth rates in 1954. Our index is essentially the crude birth rate we would expect if the Commonwealth 1954 rates were applicable to the women of each age group in each municipio and each year; consequently, the index should be positively related to the observed birth rate and have a coefficient approximately equal to one.

Death rates may also be expected to influence the relation between the level of desired family size and the birth rate necessary to achieve that level. Since some infant deaths are probably not recorded in Puerto Rico, especially in rural areas, figures on infant mortality may not be reliable in the early 1950s. Crude death rates may therefore be a better measure of survival probabilities both over time and cross-sectionally;[1] both crude and infant death rates are considered in our final analysis.[2] High death rates mean that high birth rates

[1]Earlier attempts to incorporate infant mortality rates as explanatory variables led to insignificant results.

[2]For a discussion of biological evidence for various lag structures between child death rates and birth rates see Schultz [1969a and 1969b]. A two to five year lag is considered in our model.

are required to achieve any given desired family size since fewer children may be expected to survive. Furthermore, in the long run the prevalence of high death rates will tend to be associated with earlier marriage and closer spacing of children, since the possibil of replacing a lost child late in a marriage is reduced by the lowe fecundity of the woman as well as by the greater probability that t wife or husband will have died.

FEMALE LABOR FORCE PARTICIPATION EQUATION

Fertility and female labor force participation are interrelated decisions for married women and thus jointly determined, in part, by a common set of economic forces. Overall female labor force participation depends, of course, on the joint behavior of both married and unmarried women and on the proportion of each in relation to the total number of women of working age. Our equation relates the female labor force participation rate (women in the labor force as a percentage of the number of women 14 and over) to income, the proportion of women legally married, the proportion consensually married, and a number of exogenous variables, one of which is an index of demand for female labor services. Since participation may be severely curtailed by the presence of small children in the home, the number of children under five per woman of childbearing age has also been included.

Although income is endogenous and clearly determined in part by whether or not women engage in child rearing activities or participate in the labor force or work at home production activities, income also affects the willingness of women to participate; hence, we consider it to be jointly and simultaneously determined with participation rates. The number of small children present is the result of past decisions on family size, which in turn are related to a woman's participation in the labor force, over certain portions of her life, and for certain periods of time. Thus, cross-sectional differences in this aspect of the age composition and in labor force participation rates should be related: the more small children around, the lower the female labor force participation rates. Since in general high birth rates tend to

be associated with proportionately large numbers of small children in the population, this in turn indirectly relates female labor force participation and crude birth rates.

As has been well documented, the extent of unemployment also affects the desire of a woman to work and her willingness to enter the labor force, although perhaps in different ways at different times. A high rate of unemployment may discourage a woman from seeking employment and thus, technically, she would not be counted as a member of the labor force. On the other hand, when the main family bread winner becomes unemployed for any extended period of time, other members of the family may be driven to seek supplementary employment. Cross-sectionally, differences in unemployment rates will be positively or negatively associated with variations in female labor force participation rates depending on whether the "discouraged-worker effect" or the need for supplementary income predominates. Since cross-sectional differences probably reflect persistent differences more than they do transient ones, and since the entry of women into the labor force to provide supplementary income is by nature a rather temporary affair, we might anticipate that higher than average unemployment levels would be associated with lower than average rates of female labor force participation. (See Mincer [1966].)

Because continuing in school is an alternative to early marriage, legal or consensual, we have also included a variable reflecting the proportion of children 14-17 still in school. Were the supply of schooling perfectly elastic it would be necessary to treat the variable as endogenous, since the choice between school and early marriage would have to be made simultaneously. In Puerto Rico, however, it appears that the extent of adolescent schooling is limited largely by its availability and not the demand for it; hence, we have treated the proportion of children 14-17 in school as exogenous and expect it to have a negative effect on female labor force participation.[1]

[1] We plan to test our implicit assumption that the supply of schooling is inelastic and the demand is elastic, but we have not yet found the data to confirm or reject this hypothesis that school enrollments are truly exogenous.

On an age-specific basis, marriage in any form in Puerto Rico may depress female labor force participation, as should recent child bearing, and, despite the possible association of legal marriage with a high proportion of women past the main childbearing years, we would still expect a negative effect for Puerto Rico.[1]

INCOME EQUATION

Income in this study is measured as personal income in cash or in kind in 1959 dollars per person of labor force age (15-65).[2] Income so measured clearly depends on the extent to which women work, as well as the extent to which they marry and raise children. Marriage and childbearing, other things equal, should reduce measured income, whereas increased female labor force participation should increase it. Furthermore, the more men there are relative to women, even holding female labor force participation and marriage rates constant, the higher income should be since men, on the average, tend to hold better paying jobs.

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[1]The evidence for Puerto Rico suggests that female participation falls monotonically with age after 20-24 in contrast to the United States where there is a second rise in participation rates from 35-60 (Weller [1968]).

[2]Our variable is designed to measure the average personal income in cash and kind sources per person of labor force age, that is, 15-65. The census concept of personal income includes only money income, and it is reported in terms of a median and a size distribution for persons 14 years and over by municipality. Data from a 1953 family consumption and expenditure survey of Puerto Rico are used to adjust the census figures to include nonmoney income.

Seven income-size classes are distinguished for rural and urban families in the 1953 survey. The proportion of nonmoney to money income for these fourteen groups is computed and used to inflate the census size distribution of money income. In this procedure, it is assumed that a family contains two persons with equal incomes; thus, the income classes for families are divided in half before use in the adjustment of census data for persons (with income). In the census size distributions, the mid-point of the class is assumed to be the mean income of the group, and the mean income of the open-ended income class is estimated to conform to a Pareto distribution fitted to the high income tail of the income distribution. For the procedure to estimate average income of open-ended income class, see Miller [1953 pp 24-27]. Survey data are from Income and Expenditures of Families Puerto Rico, 1953, [1967 Vol. I, Tables 21 and 57].

The association of income levels with educational attainment is well documented (Hanoch [1965]). Years of adult schooling, which we treat as exogenous in this model, should therefore be positively associated with income. Widespread unemployment should, of course, be negatively associated. Rural areas in Puerto Rico are typically a good deal poorer than urban areas, which in large part accounts for the continuing migration from the countryside. Thus, we expect to find a negative cross-sectional association between income and the percent of the civilian labor force employed in agriculture.

TWO EQUATIONS FOR MARRIAGE

Equations (4) and (5) of our model attempt to explain the percentage of women of marriageable age who are either consensually or legally married. These two equations have exactly the same form: Higher incomes should depress consensual unions relative to legal marriages since the latter are more costly than the former; at higher income levels, however, we would expect to find fewer marriages of both sorts as adolescent schooling is prolonged and the age at first marriage is increased. Since migration is highly sex selective, the proportion of men to women varies rather markedly across regions in Puerto Rico; death rates may also contribute to such variation. Thus, we have also included the ratio of men to women in the age group in which marriage, consensual or legal, is most likely to occur, 15-34.

The exogenous variables in the marriage equations are the percentage of children, 14-17, in school, an index of demand for female labor services and the unemployment rate. All of these variables might be expected to be negatively related to legal marriages with the possible exception of death rates. Schooling is an obstacle to legal marriage but less so to consensual union. Similarly, a high labor force demand for female services may weight the choice between marriage and labor force participation more heavily in favor of the latter. Again, however, the effect should be more pronounced for legal marriage and is less clear for consensual union. Unemployment should have roughly the opposite influence to income differences, although differences in

unemployment rates presumably reflect a more temporary phenomenon and fewer differences in taste than do corresponding cross-sectional differences in income.

MIGRATION EQUATIONS

Central to the dynamic workings of our model are the differentia propensities to migrate of different age and sex groups in the population. To capture the important effect of migration on the age-sex structure of the population we have introduced ten age groupings for each sex, or a total of 20 equations.

The estimation model for the 20 migration equations is derived mathematically in Appendix A as is the method of estimating interregional migration rates by age and sex cohort for the period 1950 to 1960. The model assumes that persons of a given age and sex are symmetrically attracted (or repelled) by the same weighted sum of differences between the conditions in their locality and elsewhere. An educational variable is introduced to reflect the widely observed and readily explained phenomenon that better educated persons are more responsive to migration opportunities. Much work remains to mak this initial model more realistic. Further refinements in the model are needed to test whether the levels of conditions in the origin anc destination region exert a distinguishable effect on migration rates that is independent of the difference in levels at the core of our migration model.

Data constraints are particularly severe in the migration equation. The economic incentive provided to the potential migrant by the labor market is imperfectly reflected in our estimation equation by the per capita level of personal income. Wage rates, distinguished if possible by sex, would clearly be more appropriate for our purposes. The costs of migration, on the other hand, are elusive and, at a minimum, the distances between labor markets should in the future be treated as or easily quantified component in the costs of migration. Only the distance to New York or mainland United States is explicitly considered in our model. The relative advantages of one location as opposed to

another are measured crudely in our estimation equation by income differences, differences in the unemployment rates, and differences in the level of the demand for female labor services as measured by the index described above. On the whole, it should be clear that we are least satisfied with our current treatment of migration among the many and varied elements of our model.

III. DATA AND TECHNIQUES

Observations on various demographic and economic characteristics o the population in administrative units of Puerto Rico known as municipi are available from the 1950 and 1960 U.S. censuses.[1] These data have been supplemented by additional values for the mainland United States from the Special Report on Puerto Ricans in the U.S. and by vital stati tics records of the Department of Health of the Commonwealth of Puerto Rico.[2] We make use of the data on births and deaths for the period 194 1960. Figure 1 shows the location and Table 4 the name of each municip used in our analysis.

BASIC DATA SOURCES

The results reported below for Equations (2)-(5) use a pooled sample of observations for 75 municipios in Puerto Rico for the two census years, 1950 and 1960. A dummy variable, 1 in 1960 and 0 in 1950, has been included in these analyses to allow for a change in intercept. Slopes for the two census years are assumed to be equal; however, all of the estimates have been separately computed for the two years, and, where wide divergence occurs, note is taken.[3] The migration equations, of course, involve data from both censuses, but since both are needed to compute an estimate of the migration, there is only one observation per municipio. All regressions, except the migration equations, are weighted to reflect the size of the relevant population: women over 14 for the female labor force participation and marriage equations, population 15-64 for th income equation, and total population for the birth rate equation.

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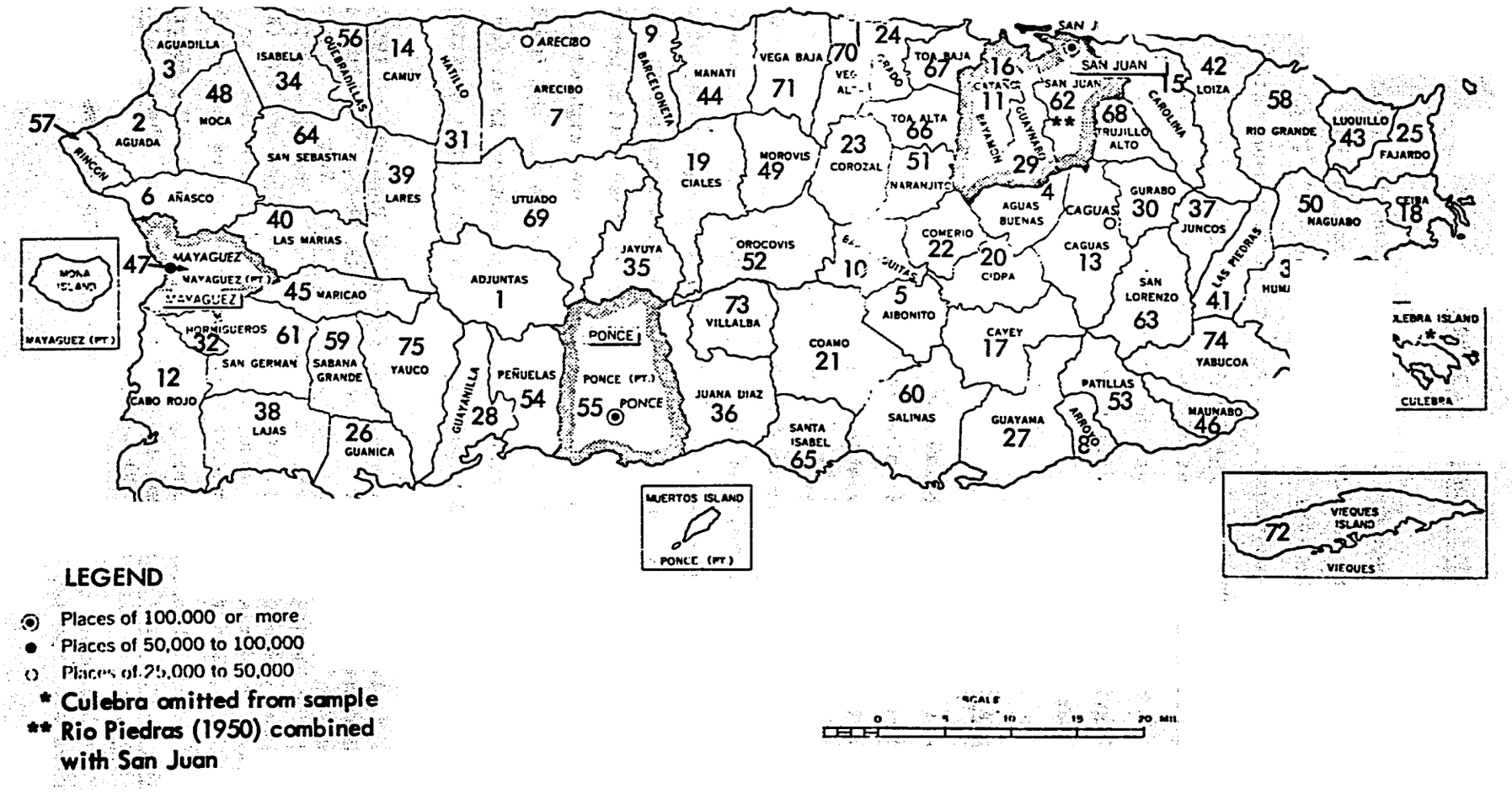
[1] U.S. Bureau of the Census, U.S. Census of Population, 1950, Vol. II, "Characteristics of the Population," Parts 51-54, "Territories and Possessions," and U.S. Census of Population, 1960, Vol. I, "Characteristics of the Population," Part 53, "Puerto Rico." The figures for one municipio (Culebra) were omitted in 1960 and in 1950, and one (Rio Piedras), bordering on San Juan, was combined with the San Juan figures in 1950 to yield totals comparable to the 1960 figures for San Juan.

[2] "Special Reports," Part 3, Chapter D, 1950, Vol. IV, Puerto Ricans in the Continental U.S.; and 1960, "Subject Reports," Puerto Ricans in the U.S., Final Report PC (2) - 1D.

[3] The results of the regressions for the 1950 and 1960 separate samples are reported in Appendix B.



LEGEND

- ◉ Places of 100,000 or more
- ● Places of 50,000 to 100,000
- ○ Places of 25,000 to 50,000
- * Culebra omitted from sample
- ** Rio Piedras (1950) combined with San Juan

Fig. 1—Puerto Rico

(Municipios, places of 25,000 inhabitants or more, and standard metropolitan statistical areas)

Table 4

LIST OF PUERTO RICAN MUNICIPALITIES[a]

1. Adjuntas
2. Aguada
3. Aguadilla
4. Aguas Buenas
5. Aibonito
6. Añasco
7. Arecibo
8. Arroyo
9. Barceloneta
10. Barranquitas
11. Bayamón
12. Cabo Rojo
13. Caguas
14. Camuy
15. Carolina
16. Catano
17. Cayey
18. Ceiba
19. Ciales
20. Cidra
21. Coamo
22. Comerío
23. Corozal
24. Dorado
25. Fajardo
26. Guánica
27. Guayama
28. Guayanill
29. Guaynabo
30. Gurabo
31. Hatillo
32. Hormigueros
33. Humacao
34. Isabela
35. Jayuya
36. Juana Díaz
37. Junc
38. Laja
39. Lares
40. Las Marías
41. Las Piedras
42. Loíza
43. Luquillo
44. Manatí
45. Maricao
46. Maunabo
47. Mayagüez
48. Moca
49. Morovis
50. Naguabo
51. Naranjito
52. Orocovis
53. Patillas
54. Peñuelas
55. Ponce
56. Quebradilla
57. Rincón
58. Río Grande
59. Sabana Gran
60. Salinas
61. San Germán
62. San Juan
63. San Lorenzo
64. San Sebasti
65. Santa Isabe
66. Toa Alta
67. Toa Baja
68. Trujillo Al
69. Utuado
70. Vega Alta
71. Vega Baj
72. Vieques
73. Villalba
74. Yabucoa
75. Yauco

Note:

[a]Culebra is omitted owing to its small size and atypical age structure; it is a small island with a total population of 573 in 1960. Rio Piedras is combined with San Juan in 1950 and 1960 for comparability with San Juan distinguished in the 1960 Census.

The equation explaining the birth rate presents a serious problem ata on birth and death rates are available continuously throughout he period between the censuses. Since the dynamic connection between he two variables is one of the most interesting and important aspects f our analysis, we cannot lightly discard the intervening nine years f data in order to utilize the available data on other variables. or would we willingly neglect these other variables in our analysis. t is natural to interpolate variables such as income, marriage rates, he age-sex composition of the population, female employment opportuni ies, and education, but is it wise; and what then will our results ean? A partial answer to these questions is provided in some recent npublished work of Grether [1969].

In an ordinary regression analysis, in which all values of the ependent variables are assumed to be present and some values of the ndependent variables may be missing (but these variables are, as sual, assumed to be uncorrelated with the disturbances), the efficiency f the estimates may be affected by how the missing observations are reated. In general, when it is known that the independent variables re very highly serially correlated, use of this serial correlation o interpolate missing values or even linear interpolation may result n substantial gains in the efficiency of estimation, particularly if he relationship that one is fitting does not fit extremely well. On he other hand, when the degree of serial correlation is low or the elationship being estimated fits well, interpolation may decrease he efficiency of estimation in comparison with discarding incomplete bservation vectors. Population and other demographic variables are ery slowly changing, hence highly serially correlated, so that Grether's esults suggest a gain in efficiency from interpolating observations n related variables, for example, the index of age composition and he index of demand for female labor services. The case for interpolating the income variable is quite weak,[1] whereas the case for

[1] It may be possible to use Commonwealth-wide figures to interpolate this variable; we shall explore this possibility in future

interpolation of the education and marriage-rate variables is stronger but perhaps less clear-cut than for the two indexes. On net balance we feel we have a case for interpolation, although not a strong one. The results presented here for Equation (1) do make use of interpolated values of the variables for the two census years; we intend to investigate further the question of whether interpolation has been helpful or not.

The method of interpolation was as follows: Let x_t be the series to be interpolated, where $t = 0$ represents 1950 and $t = 10$ represents 1960. Known values are x_0 and x_{10}. The annual compound rate of growth of x_t between 1950 and 1960 is given by r in

$$\log (1 + r) = \frac{1}{10} \log \frac{x_{10}}{x_0} \tag{3.1}$$

where logarithms are to base e. Interpolated values for the years between 1950 and 1960 are obtained by the formula:

$$x_t = (1 + r)^t x_0 \tag{3.2}$$

The series y_2, y_3, y_4, y_5, z_1, z_2, z_4, z_6, z_7, and z_8 have been interpolated directly using (3.2). The two indexes y_6 and z_3, however, interpolated by first using (3.2) to interpolate the component population or employment figures and then applying the fixed weights shown in Table 1 to these figures for each year.

We have made use of a simultaneous-equation technique (basically instrumental variables) in estimating Equations (2)-(5), but these equations are nonetheless widely familiar cross-sectional relationships We have not used simultaneous-equations techniques in the estimation of the migration equations here but will attempt to do so in the future. Equation (1) presents more interesting estimation problems since it is a relationship in a cross-section of time series. The relationship is, as we shall see, also implicitly dynamic, adding still another dimension to the problem of estimatio

ESTIMATION OF DYNAMIC RELATIONS IN A CROSS-SECTION OF TIME SERIES

Problems and methods of estimation using data on a number of individual units (firms, households, geographical areas, and so on)

over time are discussed at length in Nerlove [1971a]. We only summarize the essential nature of the problem and its proposed solution here.

Suppose we observe a large number of individuals, N, over a relatively short period of time, T. Let x_{it} be a vector of k independent exogenous variables; let y_{it} be an endogenous variable, the current value. For simplicity, suppose that the relationship to be estimated contains only one lagged value of the endogenous variable y_{it}:

$$y_{it} = \alpha y_{it-1} + x_{it}\beta + \gamma + u_{it} \quad , \tag{3.3}$$

where α is a scalar coefficient, β is a lxk vector of slopes and γ is a constant term. The disturbance u_{it} is intended to represent the net effect of numerous individually unimportant factors that have been omitted from the analysis, as well as any errors of approximation in the form of the relation. Since some of the omitted factors are likely to be rather specific to the individuals or geographical regions involved and time-invariant or only very slowly changing over time, the following model has been proposed for their stochastic structure:

$$\begin{cases} Eu_{it}\, u_{i't'} = \begin{cases} \sigma^2 = \sigma_\mu^2 + \sigma_\nu^2 & , i = i', t = t' \\ \sigma_\mu^2 & , i = i', t \neq t' \\ 0 & , \text{otherwise.} \end{cases} \\ Eu_{it} = 0, \text{ all } i \text{ and } t. \end{cases} \tag{3.4}$$

If the observations are arranged in groups corresponding to the individuals, that is, ordered by the first subscript, then by the second, for example

$$= (u_{11} \cdots u_{1T} \cdots u_{N1} \cdots u_{NT}$$

we can write the variance-covariance matrix of the disturbances a

$$Euu' = \sigma^2 \begin{bmatrix} A & 0 & \cdots & 0 \\ 0 & A & \cdots & 0 \\ \vdots & \vdots & & \vdots \\ 0 & 0 & \cdots & A \end{bmatrix} \tag{3.5}$$

where

$$A = \begin{bmatrix} 1 & \rho & \cdots & \rho \\ \rho & 1 & \cdots & \rho \\ \vdots & \vdots & & \vdots \\ \rho & \rho & \cdots & 1 \end{bmatrix}$$

The parameter $\rho = \sigma_\mu^2 / \sigma^2$ is often called the intraclass correlation coefficient in random-effects models in the analysis of variance. Our model of the disturbance in the analysis of cross-sections over time is an example of a random-effects model.

The matrix $\frac{1}{\sigma^2} Euu'$ has two distinct characteristic roots:

$$\begin{cases} \xi = [(1 - \rho) + T\rho] \\ \eta = (1 - \rho) \end{cases}.$$

It may be shown (Nerlove [1971b]) that generalized least-squares for a model with variance-covariance matrix given by (3.5) amounts to using transformed values of the variables appearing in the equation to be estimated, (3.3), rather than the original observations. If we write $\bar{y}_{i.}$ as the mean of the ith individual of y_{it}, $\bar{y}_{i.}(-1)$ as the mean of y_{it-1}, and $\bar{x}_{1.}$ as the vector of means of the exogenous variables the appropriately transformed equation is

$$\left[\frac{\bar{y}_{it} - \bar{y}_{i.}}{\sqrt{\eta}} + \frac{\bar{y}_{i.}}{\sqrt{\xi}}\right]$$

$$\left[\frac{y_{it-1} - y_{i.}(-1)}{\sqrt{\eta}} + \frac{\bar{y}_{i.}(-1)}{\sqrt{\xi}}\right] + \left[\frac{\bar{x}_{it} - \bar{x}_{i.}}{\sqrt{\eta}} + \frac{\bar{x}_{i.}}{\sqrt{\xi}}\right]$$

where

$$\left[\frac{u_{it} - u_{i.}}{\sqrt{\eta}} + \frac{u_{i.}}{\sqrt{\xi}}\right]$$

are disturbances having common variances σ^2 and zero covariances for all i and t. It is well known that estimates based on (3.7) for <u>known</u> values of ξ and η have optimal large-sample properties. It was shown in Nerlove [1971a] these estimates also have excellent small-sample properties; in effect knowing ξ and η converts a sample on N individuals over T time periods into a large sample if N is large even if T is small.

The problem is, of course, that ξ and η are not known <u>a priori</u>, but must either be estimated or assigned some arbitrary value. Setting $\rho = 0$ implies $\xi = \eta = 1$, so that (3.7) reduces to (3.3); hence, the assumption $\rho = 0$ amounts to ordinary least-squares on the "pooled" cross-section of time series. It has been shown in the Monte Carlo experiments reported in Nerlove [1971a] that ordinary least-squares yields estimates of α that are seriously biased upward and estimat of σ^2 biased downward. The standard errors obtained are consequen far too small. The bias in the coefficients β will depend upon ho the variables x_{it} are correlated with y_{it-1}; in the case considere in Nerlove [1971a], there was only one such exogenous variable and it was positively correlated with y_{it-1} so that the bias in its coefficient was toward zero

Since $\xi = (1 - \rho) + T\rho$, one can see that the second term in each of the square brackets in (3.7) vanishes asymptotically as $T \to \infty$: The resulting equations express the observations in terms o deviations from individual means. Estimating α and β from such

deviations is equivalent to estimating regressions having separate constant terms for each individual. As can be seen from (3.7), such regressions neglect the "between-individual" variation among the individual means as compared with generalized least squares. In the Monte Carlo experiments reported in Nerlove [1971a], it was shown that such neglect results in a downward bias in the estimated α and an upward bias in a derived estimate of ρ based upon the estimated residual variance plus the "variance" of the constant terms. In the case considered, a single x_{it} was positively correlated with y_{it-1} so that an upward bias in the estimated β was observed.

Several other methods of estimation, including an instrumental-variable method and a gradient maximum-likelihood procedure were examined in Nerlove [1971a]. The method we use here appeared to show the least bias, least mean square error, and greatest overall robustness against specification problems. This is a two-round procedure based upon a first-stage estimate of ρ, and thus the weights ξ and η, from a regression with deviations from individual means; the second round then yields estimates based on transformed observations, as in (3.7), using the estimated values of ξ and η to weight deviations from individual means and the individual means themselves.[1] The

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[1] Let $\bar{y}_{i.}$ be the individual means of the dependent variable, and let $\bar{x}_{i.}(k)$ be the individual means of the kth independent variable. We suppress our earlier separate notation for the lagged value of the dependent variable. Let $\bar{\bar{y}}_{..}$ and $\bar{\bar{x}}_{..}(k)$ be the grand means of y_{it} and $x_{it}(k)$, respectively, and let b_k be the estimates of the slopes in a least-squares regression of $(y_{it} - \bar{y}_{i.})$ on the variables $(x_{it}(k) - \bar{x}_{i.}(k))$. If s^2 is the sum of squared residuals from this regression, our estimate of ρ is:

(i) $$= \frac{\hat{\sigma}_\mu^2}{\hat{\sigma}_\mu^2 + s^2/N}$$

where

(ii) $$\hat{\sigma}_\mu^2 = \frac{1}{N} \sum_{i=1}^{N} \left\{ (\bar{y}_{i.} - \bar{y}_{..}) - \sum_k b_k (\bar{x}_{i.}(k) - \bar{\bar{x}}_{..}(k)) \right\}$$

method may be used with instrumental variables as well as in ordinary regression problems. Results for Equation (1) of the model are reported for both types of estimates in the next section.

IV. EMPIRICAL RESULTS

INTRODUCTION

This section presents estimates of the 25 equations that constitute the family decision making model. As stressed earlier, problems of specification and measurement are severe in some equations and less so in others. These empirical results are reported not as a final product of our research but as a tentative demonstration that the interactions among several types of family behavior are significant, both in accounting for variation in demographic and economic behavior and for understanding the correct underlying cause and effect relationships on which policy must act. Explicit treatment of these related aspects of behavior requires the use of simultaneous equations techniques and the consideration of a system of behavioral relations. To advance our limited knowledge of the interactions between man's environment and his reproductive behavior requires that we broaden our analytical framework to encompass a spectrum of family decision making activities.

Both ordinary least squares and instrumental variables have been used to estimate the coefficients of the equations. All excluded predetermined variables in the system have been used as instruments in obtaining the results labeled IV in the subsequent tables

BIRTH RATES

The results of fitting Equation (1) explaining birth rates in several different ways are reported in Table 5: Ordinary least-square (OLS) and instrumental variable (IV) estimates are computed using (1) the levels of the original variables, (2) the deviations of these variables from municipio means, and (3) the transformed variables, which are, as described above, a weighted combination of such deviations and the municipio means themselves.[1] The observations on which

[1] Regression results where the crude death rates are replaced by the infant death rates are shown in Appendix Table D-1. Although infa death rates are probably a better proxy for the child death rate impli by our model of reproductive behavior, probable errors in the measure-

Table 5

EQUATION 1: Y_1, CRUDE BIRTH RATE, 1950 - 1960
(t-ratios are reported in parentheses following each coefficient)

Type of Estimate	Variable Levels		Deviations from Municipio Means		Transformed Variables	
	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
R^2	.319		.565		.49	
F		18.72				
Endogenous variables						
y_2: Female labor force participation rat	.0863 (2.56)	.239 (.96)	.0407 (1.38)	-.502 (2.59)	.0491 (1.70)	-.303 (2.30)
y_3: Income ($100)	.294 (1.85)	3.22 (1.53)	-.478 (2.57)	1.59 (1.47)	-.338 (2.00)	-.494 (.30)
y_4: Legal marriages	397 (4.50)	-3.28 (2.94)	.854 (6.69)	2.86 (3.27)	.701 (6.50)	3.70 (3.32)
y_5: Consensual marriag	386 (4.53)	.767 (.91)	.789 (7.67)	2.74 (5.42)	.710 (7.20)	2.58 (5.21)
y_6: Index of female age composition	.0193 (.33)	-2.05 (2.59)	.142 (2.53)	-.433 (1.87)	.170 (3.16)	.384 (.94)
Exogenous variables						
z_2: Adolescent schooli	0521 (1.56)	-1.17 (3.35)	-.00151 (.34)	-.284 (2.72)	.0491 (1.70)	.0447 (.38)
z_4: Adult education	.742 (2.82)	-.216 (.15)	-.816 (3.18)	-2.93 (4.04)	-.737 (2.88)	-1.25 (1.27
$z_5(-3)$: Crude death lagged three years	.533 (4.30)	.0207 (.05)	.251 (4.01)	.178 (1.76)	.275 (4.3	.213 (2.05)
$z_5(-4)$: Crude death rate lagged four years	.207 (1.78)	-.286 (.68)	.151 (2.53)	.121 (1.20)	.173 (2.89)	.159 (1.51)
z_6: Unpaid family workers	.633 (6.44)	1.97 (4.34)	-.427 (4.97)	-.691 (4.86)	-.377 (4.35)	-.658 (4.69
Constan	2.57 (.41)	294. (3.56)	a a	a a	-3.46 (1.66)	-51.2 (2.68
			$\hat{\rho}$ = .879	$\hat{\rho}$ = .925		

Table 5 (continued)

Type of Estimate	Variable Levels OLS (7)	Variable Levels IV (8)	Deviations from Municipio Means OLS (9)	Deviations from Municipio Means IV (10)	Transformed Variable OLS (11)	Transformed Variable IV (12)
R^2	319		.562		.544	
F		5.68		4.30		27.29
Endogenous variables						
y_2: Female labor force participation rate	-.0883 (2.66)	.187 (1.25)	.0502 (1.71)	-.529 (2.92)	.0647 (2.26)	-.303 (2.48)
y_3: Income ($100)	-.310 (2.06)	-1.18 (1.60)	-.445 (2.39)	.488 (.57)	-.253 (1.51)	.714 (.95)
y_4: Legal marriages	.399 (4.54)	-2.06 (3.39)	.855 (6.68)	3.10 (3.81)	.664 (6.16)	3.02 (4.56)
y_5: Consensual marriages	.390 (4.60)	-.870 (1.72)	.842 (8.34)	2.27 (5.51)	.751 (7.64)	2.66 (5.76)
y_6: Index of female age composition						
Exogenous variables						
z_2: Adolescent schooling	.0536 (1.62)	-.361 (3.85)	.0208 (.48)	-.244 (2.54)	-.00358 (.09)	.0582 (.85)
z_4: Adult education	.733 (2.80)	1.03 (1.24)	.880 (3.43)	-2.28 (3.81)	-.790 (3.08)	1.95 (3.22)
$z_5(-3)$: Crude death rate lagged three years	.528 (4.30)	-.0970 (.39)	252 (4.01)	.226 (2.45)	.287 (4.50)	182 (1.99)
$z_5(-4)$: Crude death rate lagged four years	.202 (1.75)	-.390 (1.55)	143 (2.39)	.184 (2.05)	.173 (2.86)	120 (1.36)
z_6: Unpaid family workers	-.631 (6.44)	1.14 (5.91)	-.443 (5.17)	-.647 (4.92)	-.365 (4.54)	.644 (4.98)
Constant	2.01 (.34)	159. (4.14)	a a = .880	a a $\hat{\rho}$ = .938	-1.14 (.58)	37.4 (3.81)

Notes:

[a] Intercept suppressed.

[b] Variable excluded.

all regressions are based are weighted by the square root of municipio populations in the relevant year. Results are reported in the first half of Table 5 for analyses in which y_6, the index of female age composition, is included; the second half of the table, columns (7)-(12), shows the result of excluding y_6.

Perhaps the least satisfactory finding presented in Table 5 pertains to the coefficient of y_6, the index of female age compositi This index is essentially the crude birth rate we would expect in ea municipio on the basis of the 1954 age-specific birth rates for the Commonwealth as a whole and the age distribution of the female population in that municipio. Consequently, we migh: anticipate that the coefficient of this variable would be positive and close to one; it is however, negative in the OLS regressions, and in the IV regression based on variable levels significantly so. What accounts for this anomalous result? Further reflection suggests that although no dynamic effects (other than those introduced by lagged death rates) have been included explicitly in Equation (1), they are nonetheless present and likely to induce a strong downward bias in the coefficient of y_6: When the birth rate is high in a given municipio relative to the average, it is probable that it has been high for some time and that not all the reasons for this persistent difference will have been accounted for by the explanatory variables appearing in the equation. If the birth rate has been high relative to other municipios, then under certain relatively mild assumptions (see Keyfitz [1968, p. 186]) there will be proportionately more children in the population. Since little girls have an extremely low weight in our index, and since the population in high birth rate municipios will have proportionately fewer women in the fertile ages, it follows that the value of our inde will tend to be low for such municipios. Conversely, when the birth rate is relatively low in a municipio, the index will tend to be relatively high. Thus, the negative association found between the index of female age composition and the birth rate appears to be accounted

ment (registration) of infant deaths, as noted above, may detract from the analytical usefulness of this measure of child death rates.

for by time-persistent municipio effects that are not allowed for in the regressions on the levels of variables.[1]

When regressions are based upon weighted deviations from municipi means, the OLS estimates of the coefficient for y_6 becomes positive and significant, but the IV estimate remains negative. The estimated proportion of residual variance accounted for by time-variant municipio effects, β, is very large, 88 to 92 percent; consequently, there is little difference between the estimates based on the deviatio and the transformed variables. With the transformed variables the OLS and IV coefficients for y_6 are both positive, but still substantially less than one. Since it did not appear to be possible in this analysi to isolate in an entirely satisfactory manner the potentially importan effect of the age composition on the crude birth rate, the second half of Table 5 reports estimates for Equation (1), excluding the index of female age composition.

The coefficients of legal and consensual marriage rates are positive and approximately of the same magnitude, except for the IV estimates based on variable levels. In the final IV estimates based on the transformed data, an increase in legal marriages increases the brith rate by somewhat more than does an increase in consensual marriages, confirming that with appropriate allowances for interactions among household decisions, the secular shift from consensual to legal unions had contributed to an increase in birth rates, other things equal. The participation of women in the labor force is associated, as we anticipated, with lower birth rates, according to the IV estimates based on either the transformed data or deviations from municipio means. A reverse pattern holds for the variable levels.

In all regressions the coefficient for income was negative in the OLS regressions and positive or insignificant in the IV regressions. In a period of prosperity and Commonwealth growth, our particular

[1]This effect is particularly marked in the unweighted regressions. In this case, the coefficient of y_6 is also negative for the IV regressions. Furthermore, the adjustment described in what follows raises the value of the coefficient close to one in all instances, not merely for the IV regressions as reported below.

measure of real personal income in cash and kind per person of labor force age increased on the average 45 percent from $474 in 1949 to $687 in 1959.[1] Based on the final transformed data, the IV income coefficient is negative but insignificant (column 6), and it is positive and insignificant if the index of female age composition is excluded from the regression (column 12). This suggests the tentative conclusion that the number of children (not necessarily a quality adjusted stock) was not a superior good in Puerto Rico during this stage of economic development. Other variables that contributed to the rise in income, however, were also directly responsible for reductions in the birth rate, such as adult education.

Although we hoped to capture some elements of the costs of rearing children by means of the variable z_2, our analysis does not confirm a very strong effect; although negative signs prevail generally, only in the case of the IV estimates based on deviations from municipio means are the t ratios substantial. It is possible that more sophisticated analysis using data on factors affecting both the family demand for schooling and the community supply of school facilities would permit one to disentangle what we believe to be a potentially important determinant of desired family size.

With the exception of the regression estimates on variable levels the coefficient for adult education is consistently negative and generally significant. The magnitude of the IV estimates of the education effect on birth rates is often twice that of the OLS estimates.

In all regressions the crude death rates lagged three and four years are positively associated with current birth rates.[2] The coefficients and t ratios for the crude death rates lagged three and four years are more substantial than those estimated for death rates lagge

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[1] From 1945 to 1965 per capita national real income increased on the average about 4 percent per year.

[2] The regressions results where the crude death rates were replac by the infant death rates were less satisfactory, as seen in Table D-1. In the case of deviations from municipio means and transformed data, satisfactory OLS estimates (positive and significant) were obtained; but the IV estimates were small and generally not statistically significant.

two or five years. The three and four year lagged death rates are, therefore, adopted in the specification of Equation (1) estimated he Depending on the method of estimation, the four-year effect of a once-and-for-all unit change in the crude death rate is a .4 to .7 change in the crude birth rate. The IV estimates of the coefficient for the crude death rates are consistently smaller and less significant than the OLS estimates, and, moreover, the magnitude of these estimates is greater if the regressions are based on the variable levels rather than on the deviations or the transformed data.

Contrary to our expectations, the proportion of unpaid family workers in the labor force is negatively related to the birth rate and significantly so, except in the regressions on variable levels We conclude from this evidence that our proxy for the productive employment of children within the family has reached such a low leve in Puerto Rico by 1960 (1.25 percent) that it no longer reflects accurately the differences in child rearing and working patterns that it was intended to measure but an entirely different effect.

To trace through all of the indirect repercussions on the birth rate flowing from the other behavioral equations and implicit identi ties of the complete model requires simulation. This will be attemp in a later study to evaluate policy alternatives and their influence on the birth rate. On the whole, the final IV estimates of the birt rate equation based on the transformed variables appear to be satisfactory and to support our prior hypotheses. Possible exceptions ar the magnitude and erratic behavior of the index of female age composition, and the sign of the unpaid family worker variable. The di ficulties we had with these two variables may be related.

FEMALE LABOR FORCE PARTICIPATION RATE

The results of estimating Equation (2), explaining female labor force participation rates, are reported in Table 6. The observatior on which these regressions are based are weighted by the square root of the number of women 14 years of age and older, that is, those whc could be counted in the labor force in that municipio. Columns (1) and (2) report OLS and IV results including marriage rates; columns

Table 6

EQUATION 2: Y_2, FEMALE LABOR FORCE PARTICIPATION RATE, 1950 AND 1960 DATA COMBINED
('t ratios in parentheses)

of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.628				.626			
F			4.06				5.78	
Endogenous variables								
y_3: Income (100$)	.353	(.92)	-.618	(.26)	.353	(.93)	-.967	(.9
y_4: Legal marriages	.0413	(.15)	2.88	(1.43)	-	-	-	-
y_5: Consensual unions	-.0439	(.17)	1.25	(.85)	-	-	-	-
y_8: Children under 5 per woman 15-49	-11.6	(2.21)	-77.0	(1.76)	-12.0	(2.51)	-30.1	(1.67)
y_9: Index of age composition	2.42	(1.29)	-9.83	(.74)	-2.05	(1.32)	4.90	(1.01)
xogenous variables								
z_2: Adolescent schooling	-.269	(3.20)	.255	(.64)	-.288	(3.64)	-.145	(1.14)
z_3: Index of demand for female labor services	.223	(5.45)	.163	(1.45)	.222	(5.69)	.183	(2.81)
z_7: Unemployment	-1.48	(5.56)	-.639	(.78)	-1.52	(5.82)	-1.58	(4.73)
ummy (1960 = 1)	-2.17	(.83)	-17.3	(.95)	1.29	(.58)	6.14	(1.01)
Constant	74.8	(2.30)	116.	(.59)	69.8	(2.22)	-54.6	(.57)

Note:
- Variable not included in re ssion.

(3) and (4) report results excluding marriage rates. Only in the OLS regressions do consensual marriages have the expected negative sign, but the stronger converse results of the IV analysis cast doubt on the hypothesis that the prevalence of marriage independently depresses participation rates. The inverse relationship between the number of young children per adult woman and female labor force participation is unaltered by the exclusion of marriage rates [see columns (3) and (4)]. Income has the expected negative sign in the IV estimates, but low t ratio. The index of age composition has the expected positive effect in the IV regression without marriage rates.

Our index of the demand for female labor force services works well in explaining geographical variation in female labor force participation. The "discouraged worker" effect of unemployment clearly predominates in these results as it does in most U.S. cross-sectional studies of women's labor force participation. The effect of a high proportion of adolescents in school, however, is negative in the OLS regressions, reversing sign and losing significance in the IV regression that includes marriage rates.[1]

The effect of personal income on women's participation rate estimated in Equation (2) is a combination of the effects of the income received by men and women. If it is assumed that men's earnings are not affected by women's labor force commitments, men's earnings would exert a pure "income effect" on women's participation rate. Women's earnings, on the other hand, contain both "income and substitution effects"; the "substitution effect" must be positive and the "income effect" we expect to be negative. Because men in Puerto Rico receive

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[1]The regression results in Tables B-1 and B-2 show that the negative (gross) income effect on labor force participation of women reversed from 1950 to 1960, and young children are becoming less of a constraint on women's participation in the labor force. The effect of adolescent schooling and unemployment also diminished in 1960 compared with 1950. The index of demand for women's labor force service performs equally well in both periods. Exclusion of San Juan (Table C-1) from the combined 1950 and 1960 weighted regression indicates an increased effect of income, the age composition index, unemployment, and the demand for women's labor services on participation outside of San Juan. On the other hand, although the OLS estimates for adolescent schooling rate are similar to those including San Juan, IV estimates are less satisfactory when San Juan is eliminated.

about four times the income women received in 1959,[1] we infer that the implicit weight attached to men's income in Equation (2) is approximately four times the weight of women's income.[2] The weak negative effect of total personal income, y_3, on women's participation therefore suggests that the "income effect" is negative and exceeds in absolute value the "substitution effect" by a factor of about five.[3]

PERSONAL INCOME

Results for Equation (3) are presented in Table 7. The observations are weighted by the square root of the number of persons of

[1]Men who received income received twice the money income on the average as women who received income; twice as many men as women received money income in 1959 [U.S. Census of Population, 1960, Table 58].

[2]If we assume total personal income is the sum of male and female income:

$$y_3 = y_{3m} + y_{3f}$$

and that the appropriate form for Equation (2) is

$$y_2 = \alpha y_{3m} + \beta y_{3f} + Z$$

where Z comprises the effects of additional factors, the constant and the disturbance term, and α is the "income effect" and β is the algebraic sum of the "income and substitution effects." Then in our estimates of Equation (2),

$$y_2 = \gamma y_3 + Z$$

The parameter γ can be expressed as a weighted sum of α and β, where the weights are approximately the proportions of male and female income in the total of personal income:

$$\gamma = \alpha \frac{y_{3m}}{y_3} + \beta \frac{y_{3f}}{y_3}$$

[3]Among married women in the United States in 1950, the response of participation to a change in personal income (holding constant the proportion of income of men and women) may be estimated from some of the results of Cain's regression analysis [1966, Table 11, lines 1 and 3] as about -.12 to -.14 compared with our IV estimate for Puerto Rico of .98. Incomes were, of course, three times as large in the United States as they were in Puerto Rico. Some of the differences may occur because Cain used a single-equation model. Our OLS estimates of the income coefficient are one-fifth as large as our IV estimates. There is also evidence that the size of the negative response of participation to income is declining in Puerto Rico from the analysis of 1950 and 1960 census data separately. See Tables B-1 and B-2.

Table 7

EQUATION 3: Y_3, INCOME PER PERSON OF LABOR FORCE AGE, 1950 AND 1960 DATA COMBINED
(t-ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.945				.945			
F			176.84				31.40	
idogenous variables								
y_2: Female labor forc participation rate	-.0138	(.79)	-.278	(2.02)				
y_4: Legal marriages	.118	(2.56)	.0085	(.03)	.119	(2.58)	401	(1.60)
y_5: Consensual marriage rate	.0875	(1.74)	-.197	(.59)	.0904	(1.81)	520	(1.81)
y_8: Children under 5 per women 15-49	-1.08	(1.05)	-1.63	(.23)	-.958	(.95)	.15	(.78)
y_{10}: Sex ratio	4.15	(1.07)	2.34	(.25)	4.27	(4.05)	.2	(2.33)
xogenous variables								
z_4: Adult educati	1.09	(13.7)	.778	(2.87)	1.10	(14.1)	87	(2.62)
z_7: Unemployment	-.0856	(1.55)	-.490	(2.08)	-.0617	(1.33)	0526	(.57)
z_8: Agriculture	-.0882	(8.56)	-.184	(2.75)	-.0834	(10.0)	0618	(2.19)
ummy (1960 = 1)	.813	(3.65)	-1.66	(1.61)	.894	(4.52)	846	(.15)
onstant	-3.68	(1.39)	19.8	(1.57)	-4.60	(1.93)	.0	(1.62)

ote:
- Variable not included in reg

labor force age, from 14 to 64, in that municipio. Columns (1) and (2) show OLS and IV results for regressions including female labor force participation rates; columns (3) and (4) exclude the participation rates. The negative simple association between female labor force participation and income, which we interpreted above as the predominance of the income over the substitution effect on women's participation, is not reversed when the income equation is directly estimated; in fact, this relationship persists and is exaggerated in magnitude when the supposedly more appropriate IV method of estimati is employed.[1]

When the female labor force participation rate is excluded from the regression, the marriage rates and sex ratio bear coefficients of the anticipated sign and approximate magnitude in the IV regressions. Among the exogenous variables, adult education is associated with increased, and agricultural composition of the labor force with decreased, levels of personal income. Unemployment rates are not associated with distinctly lower personal income levels in the IV regression in which female labor force participation rate is excluded, possibly because of the difficulties of measuring unemployment comparably between modern and traditional communities, and between areas where urban and rural economic activities predominate.[2]

[1] Regressions using the 1960 data alone yield IV estimates for female labor force participation that are of the correct sign (positive) and exhibit sizable t ratios. See Table B-4.

[2] Examining the regression results for 1950 and 1960 separately in Tables B-3 and B-4, there is evidence of female participation contributing positively to personal income in 1960 as noted above. The effects of marriage, excluding female participation, increased between periods, as did the effects of the sex ratio. The effect of the child-woman ratio diminished by 1960. The strong effect of education increased somewhat by 1960, but the most notable change occurred with respect to the agricultural disequilibrium. In 1950 the association between agricultural composition and personal income is inverse but weak, yet by 1960 the magnitude of this relationship is greater and exhibits a substantial t ratio. Outside of San Juan the effect of marital status, unemployment, and agricultural composition on income is more pronounced. Of particular interest is the increased importance of sex composition and the decreased importance of education in the income equation estimated for the municipalities other than San Juan.

LEGAL AND CONSENSUAL MARRIAGES

Since the two marriage equations involve the same explanatory variables, except for the age-composition index, it is useful to con sider them together. Results are reported in Tables 8 and 9 for bot OLS and IV methods of estimation, omitting in the second set the age composition index. Observations are weighted as in Equation (2).

In the OLS regression for legal marriages, the estimated coeffi cient for the index based on the age composition of women in the municipio is as anticipated, positive but less than one. However, the sign of the age-composition index reverses sign in the IV regres sion, and the behavior of the age-composition index in Equation (5) is also erratic. We conclude, therefore, that the regional differences in age composition are not strongly associated with difference in the occurrence of the two types of marital unions, and hereafter, our interpretation focuses on the second set of regressions, omittin the age-composition index.

As one might anticipate, income appears to have a positive effe on legal marriage and a negative one on consensual marriage, whereas the reverse is true of unemployment. According to the OLS regressio the more men per woman of highly marriageable age in the municipio the more frequent are consensual marriages, while the prevalence of legal unions is not strongly affected. The IV estimates indicate th effect of the sex ratio on increasing legal marriages exceeds that on consensual unions. Greater availability of schooling for adolescents increases consensual unions relative to legal marriages, since the former may be more consistent with continued schooling, but slightly decreases the combined frequency of all forms of marriage. The index of demand for female labor services is consistently associ with fewer legal and more consensual marriages, although the t ratio for these estimates are low

On net balance, the percentage of women currently married appea to decrease with an increase in either personal income or adolescent schooling, and to increase somewhat with the local ratio of men to

Table 8

EQUATION 4: Y_4, PERCENT OF WOMEN (14+) CURRENTLY LEGALLY MARRIED, 1950 AND 1960 DATA COMBINED
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.485				.482			
F			26.09				2.80	
Endogenous variables								
y_3: Income (100 $)	.718	(2.56)	2.09	(3.19)	.866	(3.82)	.815	(2.49)
y_7: Sex ratio	-.974	(.23)	1.57	(.10)	-1.45	(.34)	19.9	(1.51)
y_{11}: Index of age composition	.466	(.89)	-3.30	(2.31)				
Exogenous variables								
z_2: Adolescent schooling	-.306	(5.45)	-.367	(4.66)	-.310	(5.53)	-.381	(5.23)
z_3: Index of demand for female labor services	-.0561	(1.85)	-.507	(.92)	-.0546	(1.80)	-.0127	(.26)
z_7: Unemployment	-.232	(1.21)	-.245	(1.00)	-.243	(1.26)	-.260	(1.15)
Dummy (1960 = 1)	4.55	(3.15)	-2.08	(.53)	3.73	(3.37)	5.69	(3.06)
Constant	37.7	(1.40)	224.	(2.78)	61.4	(12.7)	40.7	(3.04)

Note:
Variable not included in regression.

Table 9

EQUATION 5: Y_5, PERCENT OF WOMEN (14+) CURRENTLY CONSENSUALLY MARRIED, 1950 AND 1960 DATA COMBINED
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.581				.579			
F			20.77					
Endogenous variables								
y_3: Income (100 $)	-.677	(2.31)	-3.61	(2.40)	-.813	(3.70)	-1.16	(4.00)
y_7: Sex ratio	12.5	(3.02)	2.63	(.15)	12.5	(3.03)	7.74	(.65)
y_{12}: Index of age composition	-1.82	(.70)	29.7	(1.68)	-	-		
Exogenous variables								
z_2 Adolescent schooling	.232	(4.16)	.451	(3.22)	.240	(4.43)	.278	(4.22)
z_3 Index of demand for female labor services	.0173	(.59)	.0671	(1.03)	.0178	(.61)	.046	(1.05)
z_7 Unemployment	.319	(1.71)	.0287	(.09)	.315	(1.69)	.167	(.82)
Dummy (1960 = 1)	-6.18	(3.84)	9.11	(1.07)	-5.35	(4.99)	-4.7	(2.79)
Constan	9.51	(.37)	-286.	(1.70)	-7.99	(1.70)	-5.0	(.42)

Note:
- Variable not included in regression.

women. The estimates for the shift in the intercept between 1950 and 1960 for Equations (4) and (5) imply that the extent of legal marriage among adult women increased 5.7 percent and that of consensual marriage decreased 4.7 percent, which would have left the proportion married about unchanged or somewhat increased over the decade, had the other variables in the model not changed. The index of demand for female labor services and the unemployment rate do not appear to exert a substantial effect on the proportion of women currently married but they may change the composition of marriage with possible implications for other behavior, such as fertility.[1]

MIGRATION EQUATIONS

Only preliminary regression estimates of the twenty migration equations are reported in Tables 10 and 11 for men and women, each according to ten age groups. Further disaggregation of the explanatory variables is needed for specific age and sex components of the population, and interrelationships among migration patterns of various age groups should be incorporated into the model explicitly, as for example, the link between parent and child migration patterns.

The explanatory variables in each age-sex migration equation are the same: (1) population weighted mean income differences, (2) weighted differences in unemployment, (3) weighted differences in the demand for female labor services, (4) relative differences in population, (5) a

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[1]The separate regressions for 1950 and 1960 are reported in Tables B-5 through B-8. These indicate a tendency for the effect of adolescent schooling and unemployment on marriage rates to diminish between 1950 and 1960. Conversely, the sex ratio may particularly affect the frequency of consensual unions in 1960 rather than in 1950. The coefficient for income, though it does not increase as rapidly as the level of income, has a more significant association with the form of marriage in both equations in 1960 than it does in 1950. Omitting San Juan from the combined data for 1950 and 1960, the regressions reported in Tables C-3 and C-4 suggest that income, demand for female services, and unemployment exert a more pronounced effect on the composition of marriages outside of San Juan than within it whereas adolescent schooling is less important outside of San Juan. The shift in dummy variables that reflect the unexplained increase in legal marriages and the decrease in consensual unions from 1950 to 1960 are more important in San Juan than in the rest of the Commonwealth.

Table 10

EQUATIONS 6: $Y_{13}(k)$, FEMALE MIGRATION IN TEN AGE GROUPS, BETWEEN 1950 AND 1960[a]
(t ratios in parentheses)

Age Group k	R^2	Weighted Income Differences y_{15}	Weighted Differences in Unemployment Rates y_{17}	Weighted Differences in the Index of the Demand for Female Labor Services y_{19}	Relative Population Index y_{21}	"Dummy" to Account for the Extra Cost and Benefits of Moving to the U.S. y_{23}	Educational Levels z_2
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
0-4	.167	.397 (1.18)	-.150 (.39)	.0594 (.98)	-196 (1.07)	17.6 (1.07)	8.30 (.40)
5-9	.405	.875 (2.26)	.140 (.31)	.128 (1.82)	-257. (1.66)	32.1 (1.73)	-5.67 (.23)
10-14	.462	1.90 (3.43)	.298 (.48)	.140 (1.45)	-221. (1.40)	40.2 (1.54)	22.4 (.63)
15-19	.395	1.66 (3.37)	.570 (1.02)	.0606 (.68)	-245. (1.98)	57.2 (2.28)	-6.37 (.18)
20-24	.298	1.36 (3.45)	.494 (1.16)	.0286 (.41)	-256. (2.79)	63.0 (3.07)	10.3 (.35)
25-34	.192	.669 (2.36)	.209 (.68)	-.0134 (.266)	-76.6 (1.13)	17.6 (1.21)	-1.40 (.07)
35-44	.287	.346 (1.33)	.0671 (.22)	.0547 (1.12)	-241. (3.39)	49.3 (3.65)	29.8 (1.60)
45-54	.240	.178 (.62)	.00144 (.00)	.0988 (1.84)	-120 (1.82)	23.8 (1.59)	12.9 (.63)
55-64	.261	.222 (.71)	.570 (1.56)	.0804 (1.36)	-163. (1.33)	20.6 (1.30)	25.7 (1.22)
65+	.238	.797 (.90)	1.11 (1.08)	.199 (1.22)	-494 (1.00)	49.0 (1.09)	92.5 (1.62)

Note:

[illegible]vations for 75 municipios in Puerto Rico used in estimation migration variable expressed as annual rate per thousand persons.

Table 11

EQUATION 6: $Y_{14}(k)$, MALE MIGRATION IN TEN AGE GROUPS BETWEEN 1950 AND 1960[a]
(t ratios in parentheses)

Age Group k	R^2	Weighted Income Differences y_{16}	Weighted Differences in Unemployment Rates y_{18}	Weighted Differences in the Index of the Demand for Female Labor Services y_{20}	Relative Population Index y_{22}	"Dummy" to Account for the Extra Cost and Benefits of Moving to the U.S y_{24}	Educational Levels z_2
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
0-4	.085	.305	.00229	.0249	158.	13.1	4.29
		(.91)	(.01)	(.41)	(.82)	(.79)	(.21)
5-9	.621	2.52	-.247	-.00629	381.	47.7	8.61
		(7.41)	(.69)	(.11)	(2.93)	(3.25)	(.44)
10-14	.666	6.17	1.13	-.155	422.	90.1	-2.02
		(8.19)	(1.64)	(1.43)	(2.55)	(3.13)	(.01)
15-19	.519	3.79	1.05	-.0317	-396.	102.	11.8
		(5.21)	(1.53)	(.29)	(2.77)	(3.34)	(.27)
20-24	.359	1.62	-.295	.129	-139.	43.6	-30.5
		(2.76)	(.61)	(1.65)	(1.40)	(1.79)	(.87)
25-34	.132	.151	.0111	.0765	-48.6	11.3	-16.7
		(.33)	(.03)	(1.14)	(.54)	(.57)	(.59)
35-44	.146	.218	-.274	.0901	-91.2	18.3	-1.62
		(.66)	(.79)	(1.60)	(1.04)	(1.16)	(.08)
45-54	.113	.234	-.0710	.0206	-138.	21.3	17.2
		(.85)	(.23)	(.43)	(1.58)	(1.63)	(.98)
55-64	.093	.251	-.294	.0511	5.80	.152	2.88
		(.87)	(.88)	(.99)	(.04)	(.01)	(.17)
65+	.354	.105	-1.18	.526	-1309.	100.	87.4
		(.10)	(.95)	(2.85)	(1.25)	(1.81)	(1.42

Note:

[a]Observations for 75 municipios in Puerto Rico used in estimation. Migration variable expressed as annual rate per thousand persons.

variable expressing the additional costs and benefits of moving to the mainland United States, and, finally, (6) a variable reflecting differences in educational levels among municipios unweighted by population. All variables with the exception of education are weighted differences between the municipio to which the observation refers and all other municipios of Puerto Rico and the mainland U.S. Puerto Rican-born population. Although the mainland U.S. figures are used in computing the values of the explanatory variables, we do not consider a separate independent observation for in-migration to the mainland United States but fit our equation only to the 75 Puerto Rican observations.[1]

We have estimated all equations by ordinary least squares without weighting the observations; however, in a later study we intend to use instrumental variables and appropriately weight the migration rates. The age groups used in our analysis refer to ages in 1950; thus the 0-4 group was 10-14 in 1960. Although equations have been estimated for each of these age groups, we did not in fact anticipate that all equations would fit equally well; indeed, there are groups, such as children ages 0-4 or retiring adults, who tend to migrate little and, when they do, different factors from those included in our model may be responsible, such as parent migration and climate. We thus expect no direct explanatory power for our equations for such groups. To the extent that these groups are "taken along" by groups who do migrate, however, the equations may reflect certain unspecified indirect effects.

Female migration rates are generally less well explained than ma rates. The highest multiple correlations occur for males who were

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[1]See Appendix A for the practical difficulties in estimating the migration flow to the United States from Puerto Rico. Since, in principle, it should be possible to compute the migration to the United States from the other information in the model, we cannot regard the 76th migration observation as strictly independent. These migration equations are, therefore, estimated for only the 75 Puerto Rican municipio migration rates as estimated by Census-Survival techniques explained in Appendix A.

between 5 and 24 in 1950; the correlations for females are much more uniform across age groups, but higher values are also found for females who were between 5 and 19 in 1950.

Income effects on migration are generally strong and positive as expected for age groups when migration rates are high and where also the explanatory power of our equations is high. Differences in the demand for female labor services appear to explain migration, even of females, hardly at all; occasionally respectable t ratios, however, are accompanied by the expected positive signs. Relative population effects are predominantly negative. Because of the way these variables are constructed we interpret this as implying that migration occurs more than proportionately from rural areas with relatively small populations to urban areas with relatively large populations. Unemployment works poorly as an explanatory variable, often having the wrong sign and rarely exhibiting satisfactory t ratios. Education is also not a useful explanatory variable, perhaps because in this preliminary analysis it refers to the percentage of adolescents in school rather than to the education attainments of each specific age and sex cohort.

On the whole, the preliminary migration equations must be regarded as unsatisfactory, and more work will be necessary to find the most useful level of disaggregation and the appropriate specification of the model for understanding migration behavior of each of these groups in the Puerto Rican population.

V. IMPLICATIONS FOR POLICY AND FURTHER RESEARCH

FEASIBILITY AND NECESSITY OF A SIMULTANEOUS EQUATIONS MODEL

Regardless of what one may think of the individual equations making up our model of family decision making in Puerto Rico -- and we have more to say about these individual equations below -- it seems clear that it is both possible and desirable to consider a set of decisions jointly: A family's decisions with regard to its ultimate size, its location, its form, and whether the woman shall work and at what periods of her life are all jointly determined variables over the lifetimes of the individuals making up the family and thus jointly determined when aggregates of individuals are considered at a point in time.

We have seen that the crude birth rate in an area is strongly influenced by the proportion of the adult women in legal and consensual unions. The extent of female participation in the labor force contributes to somewhat lower current fertility, and conversely, women are inhibited from entering the labor force by the presence of young children and higher levels of personal income. The form and prevalence of marriage in turn is influenced by income levels and the local sex balance among those of marriageable age, which is powerfully affected by migration. These are all endogenous influences that are explained jointly by the model. Education, opportunities for women to work, the degree to which children are employed within the family, and mortality also influence births, both directly and through their influence on the other endogenous variables of the model. To understand how such variables exert their effect, one must trace both direct and indirect paths of influence.

Although form of marriage does not appear to have a significant effect on female labor force participation, income and the number of children under five do so importantly. The latter provides a strong but indirect connection between birth rates and female labor force participation. Such an association also results from the common

causal influence of the opportunities for a woman to work outside the home as measured by our index of demand for female labor services.

The strongest influences on incomes appear to be the extent to which a region is agricultural, the educational attainments of its inhabitants, and the ratio of the sexes. The latter is influenced greatly by migration and provides yet another indirect link between demographic composition and births. Since income also influences form of marriage, and this in turn does appear to influence fertility, still another chain of connection appears in the system.

Migration is affected primarily by economic differences among regions, but, because it is so age and sex selective, it is greatly influenced by, and greatly influences, demographic composition.

CLOSING AND SIMULATING THE MODEL

As indicated above, our model is not fully closed, containing many more endogenous demographic variables than are in fact "explained" by the model.[1] In principle, the age groups of each sex present in a given population are related by a series of identities. For a given area and each sex, let

$N_k(t)$ = the population in the kth age group during period t;

$D_k(t)$ = deaths of members of the kth age groups during period t;

$M_k(t)$ = net in-migration of members of the kth age group during period t.

[1] For example, the number of children under five years per adult woman in a region is a function of the number of children born in the past five years, the number who have migrated (presumably with their parents), and the number who have died, as well as a similar list of variables determining the number of adult women present. As birth rates and migration rates change in response to changes in the exogenous variables of the model, therefore, and as death rates change, both the number of children under five and the number of adult women in a region will slowly vary.

If the kth age group consists of persons whose ages vary by more than one time period (for example, a period is one year and an age group consists of persons aged 25-29 years), we need als

$A_k(t) = B_{k+1}(t)$ = the number of persons "aging out of" group k and "aging into" group k + 1 during period t.

For the first age group, births must also be considered, but these are simply those who "age into" the first age group, $B_1(t-1) = B(t)$. The basic population identities are thus

$$N_k(t) = N_k(t-1) - D_k(t-1) + B_k(t-1) - A_k(t-1) + M_k(t-1) \qquad (4.1)$$

for each age group k and each sex and region, k = 1, ..., K.

If the final age group, K, is open-ended, for example, 65+, aging out of that group is zero: $A_K(t) = 0$. If we knew the incidence of death and migration at each age within an age group then it would be easy to calculate the A_k and B_k, since births are known for each period. Without knowledge of deaths and migration at each individual age, approximate solutions for A_K and B_K exist. For example, if persons were uniformly distributed by age across an age group of n_k periods, a fraction $\frac{1}{n_k}$ of the total population group passes each period into the next age group; thus,

$$A_k(t) \cong \frac{1}{n_k} N_k(t) \qquad (4.2)$$

A life table shows, for each age group, the fraction of the population in that age group who may be expected to survive to their present age plus one year (or five years). Alternatively, the life table yiel the fraction who may be expected to die in the current period in each age group, δ_k. Approximately, therefore

$$D_k(t) = \delta_k N_k(t) \; . \qquad (4.3$$

Equation (4.3) assumes that death occurs homogeneously across ages <u>within</u> a given age group. To make this a useful approximation the fir age group should not exceed one year, since mortality in the first yea of life is many times higher than in the second, third, or fourth. Si

$$A_k(t) = B_{k+1}(t)$$

by definition, the following approximations result:

(a) Population 0-1:

$$N_1(t) = B(t) - D_1(t) \quad ; \tag{4.4}$$

(b) Population groups 1-4, 5-9, 10-14, ..., 55-64, o

k = 2, ... 10:

$$N_k(t) = N_k(t-1) - \delta_k N_k(t-1) + \frac{1}{n_{k-1}} N_{k-1}(t-1)$$

$$- \frac{1}{n_k} N_k(t-1) + M_k(t-1); \tag{4.5}$$

and

(c) opulation group 65+, k = K = 11:

$$_k(t) = N_K(t-1) - \delta_K N_K(t-1) + \frac{1}{n_{k-1}} N_{K-1}(t-1) \tag{4.6}$$

$$+ M_K(t-1)$$

We do not have a separate life table for each municipio. We have .ife tables for Puerto Rico in 1950 and 1960, and both crude and infant leath rates by municipio annually. Such death rates may be used to :ompute estimates of the δ_k for each municipio based on the Commonwealth-ride life tables and the municipio's age composition.

Once the model has been closed by "identities" such as those .llustrated, it will be possible to simulate the model in order to ıssess the effects of changes in the time paths of the exogenous vari-ıbles in the model. These variables are educational ones, a variable :eferring to industrial structure, death rates, the unemployment rate, :he extent to which the community is agricultural, and the extent to /hich unpaid family workers form a part of the labor force. Policy :an affect nearly all of these in one way or another, and thus the ıehavior over time of the endogenous variables of the model. Migration, /hich is endogenous in the model, will, in the long run, affect the

extent to which the community as a whole is agricultural, although this is ignored in the present formulation. Although the full implications of our model for policy can only be explored by proper simulation, some short-run qualitative implications may be drawn from the model in its incomplete form.

SOME QUALITATIVE IMPLICATIONS FOR POLICY

Variables that may be affected by public policies include the provision of educational facilities, the general level of economic activity and the degree of unemployment, the location of industry and the extent and nature of employment opportunities for women, and the provision of public health services. Birth control information may also be disseminated and such information may affect birth rates. Unfortunately, we have not been able to incorporate in our model a variable, or variables, reflecting the diffusion of contraceptive techniques throughout the population, and, consequently, our results cannot at this stage shed any light on this aspect of public policy and demographic change.

The provision of educational services affects the endogenous variables of the model and, particularly, birth rates in several ways: Two educational variables enter the birth rate equation itself; one attempts to measure the short-term effects of increased availability of schooling on parents' desires for children, the other tries to capture longer-term effects of increased education of the parents on their family-size preferences. Although we have not been highly successful in isolating these direct educational effects, there is the possibility that more refined analysis may uncover both relationships we believe to be present. In addition to such direct educational effects, however, there are indirect routes by which increased availability of educational facilities may affect fertility, and we have found evidence of the strength of these. Adolescent schooling is negatively associated with legal marriage and positively associated with consensual union; but the net effect will be a significant reduction in birth rates due to increasing adolescent education. We have

not been able to find evidence of yet another route of indirect influence via the effect of education on migration.

The general level of economic activity and its distribution geographically is captured in our model primarily by the rate of unemployment. One can think of other ways in which general activity could operate, but these are not well described by the model. Unemployment does not directly affect births in our model, but influences female labor participation, income, and form of marriage. Through each of these intervening variables, lower unemployment rates should tend to be associated with slightly increased birth rates. However, additional longer term effects may occur through the influence of unemployment rates on migration and the changes in the age-sex composition of the population resulting therefrom.

One influential economic variable in the model is our index of demand for female labor services. The existence of good employment opportunities for women outside of the home appears to have a direct positive effect on female participation in the labor force and negative effect on marriage that should tend to lower fertility slightly.

An important policy variable affecting births directly, however, appears to be the provision of public health and medical services. Lower death rates are markedly associated with lower fertility, although not necessarily with a slower rate of population growth. More research is needed to identify the contribution of public policy to these differences in death rates.

Policies that affect the attractiveness of migration and the rural-urban distribution of population may also importantly affect fertility. Although few such policy influences can be discerned from our migration equations in their present form, the strong influences of the age composition of an area on fertility suggest a potential route of indirect influence, if not one on which national policy might rely.

IMPROVING INDIVIDUAL EQUATIONS

The statistical results of the present investigation can hardly be regarded as satisfactory; their usefulness lies largely in suggesting directions for further research and the feasibility of a simultaneous-equations model of the type we have described.

Some of the least satisfactory equations are those related to migration. It may not be reasonable to attempt to explain migration for as fine an age-sex breakdown as proposed here. For example, children who were under five in 1950 and, therefore, under fifteen in 1960 cannot be expected to migrate independently of their parents. Dennis De Tray has done some preliminary calculations that suggest the feasibility of relating child migration to the migration of adult groups more explicitly. For example, in a regression relating the migration rate of children (y: males and females aged 0-4 years in 1950) to the migration rates of women grouped by age (x_1: aged 15-19; x_2: aged 20-24; x_3: aged 25-34, all in 1950), the number of children per woman aged 15-34 (x_4) and the distance of the center of the municipio in question to San Juan in hundreds of air miles (x_5), De Tray found:

$$y = .0048 - .314x_1 + 1.03x_2 + 1.67x_3 - .0017x_4 - .00063x_5, \quad (4.7)$$

$$\quad (5.80) \quad (1.22) \quad (3.20) \quad (5.51) \quad (4.43) \quad (1.34)$$

$$R^2 = 0.81$$

The numbers in parentheses below the coefficients are t ratios, and the regression was estimated by ordinary least squares. These preliminary results obtained by De Tray suggest that relating certain age-sex categories of migration to others and the latter to economic variables may be more fruitful than our present approach.[1]

It is also possible that broader aggregates of migrants will prove easier to explain than the groupings used in this study. If

[1] Dropping x_1 from the regression did not substantially change the remaining coefficients nor the overall explanatory power of the equati

broader groupings are chosen, it will be necessary to relate the finer groups to the broader ones since the former play an important role in the other equations.

One of the most serious defects in the present model is the assumption that some labor market phenomena and child school attendance rates are determined independently of the endogenous variables in the model. Unemployment is certainly not independent of female labor force participation nor, in the long run, is our index of demand for female labor services. Although any attempt to treat industrial structure that underlies our index of demand for female labor services as endogenous would surely lead to great complications, the same does not appear to be true of unemployment or child labor force participation rates. It would clearly be desirable to specify, identify, and estimate simultaneously the supply and desired demand schedules for female and child labor and thus improve on the existing form of studies of labor force participation that combine demand and supply factors in the same ad hoc regression equation.[1] To progress in future work and expand our model of the labor market sector will require at a minimum sex- and possibly age-specific wage data by region. These wage data are not now available to us for Puerto Rico.

In low income countries, particularly in rural areas, the opportunity costs of sending a child, especially an older child, to school may be very great. Under these circumstances it is difficult to maintain the assumption that the supply of places in the school system entirely determines the number of children enrolled. With data recently acquired, it may be possible to estimate the interaction of government-supply of and parent demand for education and their joint determination of the time children spend in school. Estimation of the influence that government policy and resources exert on school attendance rates, or conversely school dropout rates, is a priority objective for our further research.

[1]The difficulty of this task is testified to by Cain's unsuccessful ttempt using U.S. data [1966, Appendix H].

CONCLUSION

This Memorandum is a progress report on a continuing investigation of the economic determinants of fertility, marriage, migration, and labor force participation using simultaneous-equations econometric techniques based on cross-sectional data over time. Formulation and estimation of a simultaneous-equations model describing the determination of what are essentially jointly determined variables has two main advantages over an equation-by-equation or decision-by-decision approach. First, it permits the use of more appropriate techniques of estimation than are possible in single-equation models. Second, at least for a complete model closed by the necessary demographic identities, it permits dynamic simulation of the demographic economic characteristics of an area over time.

More appropriate estimation techniques should and do lead in almost all cases to different estimates of the significance and importance of various factors affecting behavior. For example, many studies have found a negative association between income and fertility based on cross-sectional data, but we have shown this relationship dissolves when the statistical analysis of Puerto Rican data takes account of both the simultaneous nature of the system in which the birth rate equation is embedded and the special nature of the cross-sectional data over time. The pronounced effects of both legal and consensual unions on birth rates are magnified when the simultaneous character of the model is recognized and the appropriate estimation techniques employed. Although not every estimate obtained by the techniques outlined here is more plausible in size or in sign than those obtained by ordinary regression, the estimates on the whole appear improved over those obtained by methods that treat each relationship in isolation.

The results presented here are preliminary and incomplete. But we hope that they may serve as a first step toward our ultimate goal of a complete and comprehensive structure of the behavior of the family unit incorporating demographic and labor market variables endogenously. This final model must be capable of dynamic simulations

designed to reveal the long term demographic and economic effects of policies that affect the fertility and mortality and improve the educational and economic opportunities of the evolving population.

Appendix A

MIGRATION

THE MODEL

Let p_{ij} be the probability that an individual in a given age group will migrate from i to j; p_{ji} is the corresponding probability that an individual will migrate in the reverse direction. If n_i is the total population in the ith area in the appropriate age-sex group, then $n_i p_{ij}$ is the expected number in that group who will migrate from i to j. Similarly $n_j p_{ji}$ is the expected number who will migrate from j to i, so that the expected net migration from i to j is just $n_i p_{ij} - n_j p_{ji}$ for the age-sex group in question. Th expected flow of migrants into a region i is the sum over all regions k of the net migration from k to i; thus

$$m_i = \sum_{k=1}^{K} \left(n_k p_{ki} - n_i p_{ik} \right) + b n_i z_{2i} ,$$

where K is the total number of regions. (Note that the term of the sum for k = i is $n_i p_{ii} - n_i p_{ii} = 0$.) We have also included a term reflecting the current educational attainments of individuals in municipio i, bz_{2i}, since we expect greater educational attainment will increase the propensity to migrate, at least outwardly, and should thus have a negative association with net migration into any region.[1]

Since the probability of moving between any two regions is a number between zero and one, it would be inappropriate to assume : to be a linear function of any determing variables; nonetheless, i

[1] The use of z_2, adolescents in school, rather than a direct r sure of the educational attainment of the population currently livin in the municipio, such as z_4, is designed to circumvent the obvious difficulty that results from the effect of migration: Since the better educated tend to migrate more, there will be fewer of them le in a municipio having a net outflow of population. Large numbers adolescents in school, however, suggest a high level of educationa attainment in the area without itself being directly influenced by the migration it is supposed to explain.

will assume linearity as an approximation over the fairly narrow range of variations observed in migration rates. The gains from moving from i to j must be set against costs. If the advantages of i are measured by x_i, the advantages of j by x_j, and the cost of moving from i to j by c_{ij}, we have

$$p_{ij} = a_0 + a_1(x_j - x_i) + a_2 c_{ij} \tag{A.2}$$

Substituting in (A.1) and assuming that the cost of moving from i to j is the same as moving from j to i, $c_{ij} = c_{ji}$, we obtain expected net in-migration for region i as

$$M_i = a_0 \sum_k (n_k - n_i) + a_1 \sum_k (n_k + n_i)(x_i - x_k) \tag{A.3}$$
$$+ a_2 \sum_k (n_k - n_i)\, c_{ik} + bn_i z_2$$

If the cost of moving between any two points is approximately constant, the final term may be absorbed in the first. Expressing migration as a rate per capita of the relevant population group, we find that expected net in-migration rate into i is a function of two types of variables: Proportionate population differences and weighted differences of variables measuring [illegible] ving into i

$$\frac{M_i}{n_i} = a_0 \sum_k \left(\frac{k \quad i}{n_i}\right) + a_1 \sum_k \left(\frac{k \quad i}{n_i}\right)(x_i - x_k) + bz_{2i} \,. \tag{A.4}$$

In using this result for estimation purposes a number of complications arise: Actual migration cannot be measured but only inferred by observing the resident population and its age-sex composition at two points in time (census years) and applying estimates of age-specific death rates in order to compute what the population in various age-sex groups ought to have been in the second period in the absence of any migration. The details of this computation are given in the next section of this Appendix. We use estimates of the actual in-migration rates to represent the expected in-migration rate appearin

on the left-hand sides of our 20 migration equations. We have expressed migration per capita at an annual compound rate rather than a decadal one, although the two census years, between which we measure migration, are ten years apart.

As indicated above, we have assumed the costs of moving to be the same between any two points, except, however, between any point and the mainland United States; we modify this assumption by introducing a dummy variable as follows: Suppose that moving costs are c between any two points not including the mainland United States, and c + c' between any point and the mainland United States; then the last term in (A.3) becomes

$$a_2 \sum_k (n_k - n_i)c_{ik} = a_2 c \sum_k (n_k - n_i) + a_2 c' (n_{76} - n_i) \qquad (A.5)$$

where observation 76 represents the mainland United States. Consequently each equation should contain an additional variable for each age-sex group,

$$\left(\frac{n_{76} - n_i}{n_i}\right)$$

which takes account of the greater cost entailed in leaving the island for the mainland United States as well as possible unspecified benefits. This is the second modification.

Finally, it is necessary to consider what variables reflect the relative advantages of one location as opposed to another. The variables we have chosen to examine are income differences, differences in the unemployment rates, and differences in the level of the demand for female labor services as measured by the index described above. These variables are all expressed as weighted sums of the differences between the value of the variable for other areas and the value for the area of observation. Thus, we expect that high values of the income and demand variables (meaning that other areas compare favorably on the whole with the area in question) should tend to be

associated with low values of net in-migration, while high values of the unemployment variable (meaning other areas compare unfavorably) should be associated with high in-migration rates.

MEASUREMENT OF MIGRATION

A regional migration rate is designed to express the relation between a migration flow and the population sustaining or being subjected to that flow. For convenience the migration rate is defined as the ratio of the net migration flow to the domestic population, regardless of the direction of the net movement of persons. There may appear to be no measurement problem here, but since observations of the population on which migration rate estimates are based occur at discrete intervals rather than continuously, various changes occurring to the population in that interval must also be incorporated into the analysis.

Three processes may affect the size of an open population: some persons may be born, some may die, and some may migrate into or out of the regional population. No estimates are derived or analyzed here for the cohort born between the two census dates. Migration of age cohorts completely enumerated in both censuses, however, are affected by deaths hich reduce the population size over time, or the denominator, and thus he migration ratio. Were the migration rate measured instantaneously, he frequency of death would not affect population size and thus would ot bias a direct estimate of migration. But, as stated above, measure- ent of migration based on census data must account for the effect of eath on population size.

Consider a cohort of individuals of a specific age and sex. For country experiencing no net migration, let this cohort consist of A_j persons age x at time 0, and t years later let there remain B_j ersons in this cohort of age x+t. A survival rate R_j is defined as B_j/A_j where the subscript j stands for the specific age-sex cohort. By assumption the difference between A and B is attributed to deaths. If the frequency of deaths were constant over the time interval, the

effect of death may be represented as an exponential rate of decline of the population:

$$A_j e^{-d_j t} = B_j \qquad (A.6)$$

where e is the base of natural logarithms, d_j is the instantaneous death rate, and t is elapsed time.

In an open (regional) population subjected to net migration, the difference between the base and final period population size is due to the effects of both death and migration. Assuming that the relative frequency of death and migration is constant over the interval considered, both exponential effects on population size may be written:

$$A_{ij} e^{(m_{ij}-d_{ij})t} = B_{ij} , \qquad (A.7)$$

where the i subscript distinguishes the region and m_{ij} is the age-sex-region specific average annual migration rate, greater than zero if there is net in-migration and less than zero if there is net out-migration.

Since age-sex specific death rates are not known for reg: populations in Puerto Rico, one must assume that $d_{ij} = d_j$ for all regions $i = 1,2,\ldots,N$. It is then possible to estimate m_{ij} from national survival data and regional age-sex population data from two censuses:[1]

$$m_{ij} = (\log B_{ij} - \log A_{ij} - \log R_j)/t \qquad (A.8)$$

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[1] From equation (A.7), dividing by A_{ij} and taking natural logarithms yields

$$(m_{ij} - d_j)t = \log B_{ij} - \log A_{ij} \qquad (a)$$

and since from equation (A.6), $d_j t$ is equal to $\log (A_j/B_j)$ ıation (a) is rewritten as

$$m_{ij} = (\log B_{ij} - \log A_{ij} - \log B_j + \log A_j)/t \qquad (b)$$

which is identical to equation (A.8) when B_j/A_j is replaced by R_j.

This procedure for measuring migration rests on two assumptions: (1) Censuses enumerate all (or a uniform proportion of) people and accurately report their age and sex, and (2) age and sex specific death rates are uniform across regions analyzed.

Age is often systematically misreported, contributing to "age heaping" in the age profile of the population about attractive or "round" ages. In Latin America, and in Puerto Rico in particular, age heaping does not appear to distort the census data seriously for our purposes.

A general problem of systematic underreporting arises, however, in the census enumeration of U.S. residents born in Puerto Rico. The distinction of being born in Puerto Rico may be associated in the minds of many persons with lower-class status in the United States, and therefore is likely to be subject to a downward response bias. Moreover, the 1960 census tended to underenumerate urban poor and Negro elements in the population, and probably did the same for the largely urban, relatively poor, Puerto Rican born U.S. residents. Hence, the survival technique for estimating migration from Puerto Rico to the U.S. mainland is likely to understate the net influx. For these reasons the migration equations are estimated, excluding the U.S. observation. In fact, the estimated rates of migration to the United States shown in Table A-6 imply a net inflow of 320,000 persons over the age of ten by 1960, whereas other methods of estimation suggest total migration for all ages may have been between 400,000 and 430,000. Even these latter estimates may be deficient according to Hernandez [1967], who has estimated the gross return migration flow during the 1950s as between 50,000 and 75,000. The gross inflow from Puerto Rico to the United States was therefore of the order of half a million in this decade from an initial Commonwealth population of 2.2 million to an initial U.S. resident Puerto Rican born population of 226,000 in 1950.

The second assumption underlying the measurement of migration is that death rates by age and sex are the same in all regions of the country. It is probable that death rates are higher for populations

living in rural hinterlands of Puerto Rico than for those in urban centers. It may be supposed, therefore, that estimated out-migration from remote rural areas is overstated because of higher death rates prevailing in these areas, and conversely out-migration for more prosperous urban oriented areas is understated (or in-migration is overstated). Hence, the lack of information on regional death rates may exaggerate differences in migration between rural and urban regions, though the actual magnitude of this bias appears small in comparison with the size of net migration rates in this period. Substantial differences in the incidence of mortality could conceivably lead to differences in annual migration estimates of only about .1 to .2 percent per year for the younger and older cohorts, and somewhat less for those in the prime of life. In comparison with average annual net out-migration rates of between 1 and 7 percent per year (Table A-1), this source of bias in our estimates may be regarded as negligible.

EMPIRICAL MATERIALS AND PROCEDURES

Life tables have been published for Puerto Rico for men and women in 1959-1961 and for women in 1949-1951 in five year intervals.[1] It is assumed that the ratio between survival rates for men and women in 1960 obtain also for 1950 to derive estimates of mortality by age and sex at both ends of the decade. A product of the 1950 and 1960 survival rates (for five year periods) is used as the estimated survival rate for the decade between the censuses. For example, the five year survival rate in 1950 for the cohort age x to x+4 is multiplied times the five year survival rate for the cohort age x+5 to x+9 in 1960 to obtain an estimate of the survival rate over the entire decade for the cohort aged x to x+4 in 1950 that had obtained the age of x+10 to x+14 by 1960.[2] These estimated survival rates are shown in Table A-2.

[1] 1950 Female Life Table from Carballo (1964), Table 9, p. 19, and the 1960 survivor rates from UN Demographic Yearbook 1966 (1967), Table 23, pp. 608-609.

[2] For the population over 64 in 1950 and over 74 in 1960 the age breakdown was not available by municipality. It was therefore assumed

Table A-1

ANNUAL NET IN-MIGRATION RATES, MEANS AND STANDARD DEVIATIONS FOR 75 MUNICIPALITIES OF PUERTO RICO AND NET MIGRATION RATES TO THE MAINLAND UNITED STATES
(percent)

Age of Cohort in 1950	75 Municipalities Puerto Rico				Mainland U.S.	
	Male		Female			
	Mean	Standard Deviation	Mean	Standard Deviation	Male Mean	Femal Mear
0-4	-.91	.98	-1.07	1.03	19.83	20.1
5-9	-3.16	1.41	-3.19	1.44	14.34	14.7
10-14	-7.17	3.06	-5.15	2.20	15.76	17.0
15-19	-7.28	2.78	-5.16	2.07	16.47	14.3
20-24	-5.28	1.93	-4.50	1.60	9.32	7.6
25-34	-2.22	1.34	-1.73	1.03	6.93	5.2
35-44	-1.61	1.04	-2.20	1.00	3.70	3.1
45-54	-.78	.84	-1.12	1.06	3.07	4.1
55-64	.02	.83	-.62	1.12	4.09	4.6
65 or over	.43	3.48	.86	2.99	3.61	2.0

Table A-2

SURVIVAL RATES FOR PUERTO RICO, 1950-1960

Age Cohort in 1950 and 1960	Men	Women
0-4 to 10-14	.9475	.9498
5-9 to 15-19	.9839	.9871
10-14 to 20-24	.9854	.9910
15-19 to 25-29	.9714	.9831
20-24 to 30-34	.9551	.9735
25-29 to 35-39	.9506	.9686
30-34 to 40-44	.9408	.9622
35-39 to 45-49	.9339	9535
40-44 to 50-54	9237	9472
45-49 to 55-59	8934	9316
50-54 to 60-64	8632	9102
55-59 to 65-69	8218	8781
60-64 to 70-74	7501	8220
65-69 to 75-79	6599	7469
70-74 to 80-84	5567	6292
75-79 to 85-89	4343	4889
80+ to 90+	3120	3810

The values of the environmental determinants of migration, that is, income, unemployment, the index of demand for female labor services, and population, which are used to construct variables y_{15} through y_{24} are municipal averages of 1950 and 1960 census observations. The school enrollment rates are also averages of 1950 and 1960 values within Puerto Rico.

Several gaps in matching data on the Puerto Rican born U.S. resident population require compromises in our data and preclude further analysis of actual migration from Puerto Rico to the United States. The industrial distribution of employment for 1950 is not reported for the Puerto Rican born U.S. residents, and in 1960 the distribution is reported, but only for 14 industries rather than the 21 used to define the index of demand for female labor services within the Commonwealth. Reaggregating the Puerto Rican female employment weights for the 14 available industries, an index of demand for female services of 137.13 is derived, estimated for the U.S. resident population. The school enrollment rate is reported in the 1960 census, but must be interpolated for 1950 between different age groups, to yield an estimate that 68 percent of the Puerto Rican born, 14 to 17 age cohort were in school during the 1950s in the United States. Personal income per person aged 15-64 in 1959 dollars is estimated as $1253 in 1949 and $1750 in 1959. The period average is thus more than twice the unweighted municipal average in the Commonwealth of $655. No adjustment is attempted to account for differences in the level of the cost of living between the United States and various parts of the Commonwealth of Puerto Rico.

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that in each region these four oldest age cohorts were fixed proportions of the total number over 64 and 74, respectively, as enumerated in the Census for the Commonwealth as a whole. The net migration flows were then aggregated from the 17 age cohorts shown in Table A-2 into the 10 age cohorts analyzed in this study, and shown in Table A-1.

Appendix B

REGRESSION RESULTS FOR 1950 AND 1960 SEPARATELY, EQUATIONS 2 THROUGH 5

Table B-1

EQUATION 2: Y_2, FEMALE LABOR FORCE PARTICIPATION RATE, 1950 DATA
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.684				.670			
F			4.79				9.00	
Endogenous variables								
y_3: Income (100 $)	-.883	(1.02)	3.33	(1.02)	-1.05	(1.22)	4.65	(1.63)
y_4: Legal marriage	.717	(1.50)	1.12	(.79)	-	-	-	-
y_5: Consensual marriage	.502	(1.14)	.752	(.66)	-	-		
y_8: Children under 5 per woman 15-49	-26.4	(1.98)	-62.8	(1.27)	-17.6	(1.67)	-47.4	(1.01)
y_9: Index of age composition	12.8	(4.15)	23.4	(2.58)	-10.6	(3.78)	-24.7	(2.67)
Exogenous variables								
z_2: Adolescent schooling	-.527	(4.31)	-.539	(1.36)	-.595	(5.25)	-.733	(2.95)
z_3: Index of demand for female labor services	.279	(3.81)	.0753	(.30)	.264	(3.61)	.0306	(.12)
z_7: Unemployment rate	-1.36	(3.80)	-.893	(1.62)	-1.57	(4.62)	-1.11	(2.20)
Constant	265.	(4.60)	494.	(2.49)	257.	(4.46)	575.	(3.19)

Note:
- Variable not included in regression.

Table B-2

EQUATION 2: Y_2 FEMALE LABOR FORCE PARTICIPATION RATE, 1960 DAT
(t ratios in parentheses)

Types of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.816				.798			
F			98.38				8.1	
dogenous variable								
y_3: Income (100	-.761	(2.11)	-.154	(.12)	-.775	(2.08)	-.919	
y_4: Legal marriage	-.572	(2.13)	.782	(.68)				
y_5: Consensual marriage	-.266	(1.00)	.844	(.86)				
y_8: Children under 5 per woman 15-49	-10.4	(2.71)	-8.14	(.82)	-11.5	(3.02)	-5.30	(.67)
y_9: Index of age composition	5.26	(3.08)	.427	(.06)	2.40	(1.82)	4.18	(1.69)
ogenous variables								
z_2: Adolescent schooling	-.0360	(.43)	-.0231	(.19)	-.00388	(.05)	.010	(.11)
z_3: Index of demand for female labor services	.196	(3.57)	.271	(3.26)	.255	(5.08)	.271	(3.88)
z_7: Unemployment rate	-.596	(1.92)	-.829	(1.97)	-.653	(2.15)	-.723	(2.10)
nstant	-54.7	(2.11)	-45.8	(.61)	-33.8	(1.34)	-73.8	(1.59)

te:
- Variable not included in regression.

Table B-3

EQUATION 3: Y_3, INCOME PER PERSON OF LABOR FORCE AGE, 1950 DATA
(t ratios in parentheses)

of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.926				.925			
F			38.16				24.80	
Endogenous variables								
y_2: Female labor force participation rate	.0206	(1.02)	-.140	(2.45)	-	-		
y_4: Legal marriage	.165	(2.54)	.0445	(.21)	.166	(2.56)	-.0544	(.28)
y_5: Consensual marriage	.141	(2.12)	-.0719	(.33)	.146	(2.21)	-.0844	(.43)
y_8: Children under 5 per woman 15-49	-6.13	(2.85)	-11.5	(1.58)	-6.18	(2.88)	-12.5	(1.90)
y_{10}: Sex ratio	4.95	(3.04)	5.69	(1.32)	5.37	(3.41)	7.21	(1.87)
Exogenous variables								
z_4: Adult education	.924	(5.94)	.708	(2.94)	.950	(6.19)	.896	(4.33)
z_7: Unemployment	-.0212	(.29)	-.244	(1.69)	.0238	(.42)	.0075	(.08)
z_8: Agriculture	-.0662	(4.08)	-.0872	(1.82)	-.0570	(4.21)	-.0145	(.43)
Constant	-3.89	(1.05)	12.9	(1.17)	-5.50	(1.65)	8.93	(.90)

Note:
- Variable not included in regression.

Table B-4

EQUATION 3: Y_3, INCOME PER PERSON OF LABOR FORCE AGE, 1960 DATA
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.947				.947			
F			148.36				61.6	
Endogenous variables								
y_2: Female labor force participation rate	.0248	(.71)	.293	(1.97)	-	-		
y_4: Legal marriage	.168	(2.34)	-.186	(.64)	.162	(2.28)	-.208	(.83)
y_5: Consensual marriage	.135	(1.51)	-.255	(.65)	.133	(1.50)	-.285	(.85)
y_8: Children under 5 per woman 15-49	.0686	(.06)	6.02	(1.28)	-.183	(.15)	5.21	(1.31)
y_{10}: Sex ratio	2.81	(1.80)	11.5	(1.64)	2.70	(1.74)	12.9	(2.16)
Exogenous variables								
z_4: Adult educatio	1.09	(9.39)	.959	(3.60)	1.09	(9.42)	1.01	(4.49)
z_7: Unemployment	-.0618	(.64)	.0281	(.15)	-.0766	(.81)	-.171	(1.23)
z_8: Agriculture	-.0922	(6.12)	-.0302	(.63)	-.0984	(8.00)	-.112	(5.73)
Constant	-5.97	(1.47)	-4.73	(.36)	-4.70	(1.29)	4.61	(.44)

Note:
- Variable not included in regression.

Table B-5

EQUATION 4: Y_4, PERCENT OF WOMEN (14+) CURRENTLY LEGALLY MARRIED, 1950 DATA
(t ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
	.488				.482			
			11.82				10.34	
Endogenous variables								
y_3: Income (100 $)	44	(2.44)	1.85	(1.54)	1.37	(2.36)	1.99	(1.62)
y_7: Sex ratio	8	(1.64)	-2.23	(.10)	3.8	(1.64)	7.58	(.35)
y_{11}: Index of age composition	698	(.90)	2.44	(1.63)				
Exogenous variables								
z_2: Adolescent schooling	-.316	(3.69)	-.398	(3.21)	-.303	(3.60)	-.398	(3.14)
z_3: Index of demand for female labor servic	-.0786	(1.35)	-.0469	(.38)	-.0904	(1.59)	-.096	(.79)
z_7: Unemployment	-.411	(1.63)	-.278	(1.00)	-.436	(1.74)	-.289	(1.02)
Constant	109.	(2.76)	85.	(2.23)	4.3	(8.61)	54.5	(2.57)

Note:
– Variable not included in regression.

Table B-6

QUATION 4: Y_4, PERCENT OF WOMEN (14+) CURRENTLY LEGALLY MARRIED, 1960 DATA
(t ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.421				.283			
F			1.91				1.73	
Endogenous variables								
y_3: Income (100 $)	.327	(.97)	.041	(.06)	1.08	(3.44)	1.14	(2.52)
y_7: Sex ratio	4.93	(1.09)	9.44	(.78)	3.05	(.61)	-2.37	(.20)
y_{11}: Index of age composition	2.84	(4.03)	2.98	(2.12)				
Exogenous variables								
z_2: Adolescent schooling	-.138	(1.81)	-.134	(1.67)	-.189	(2.28)	-.184	(2.18)
z_3: Index of demand for female labor services	-.181	(3.75)	-.146	(2.23)	-.129	(2.51)	-.142	(1.97)
z_7: Unemployment	.215	(.74)	.083	(.24)	-.056	(.18)	.031	(.08)
Constant	-79.8	(2.30)	-91.3	(1.24)	58.4	(10.3)	63.3	(5.51)

Note:

Table B-7

EQUATION 5: Y_5, PERCENT OF WOMEN (14+) CURRENTLY CONSENSUALLY MARRIED, 1950 DATA
(t ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.340				.309			
F			6.15				3.65	
Endogenous variables								
y_3: Income (100	----	(1.99)	-3.54	(2.20)	-.844	(1.33)	-1.67	(1.27)
y_7: Sex ratio	20.6	(2.30)	21.1	(.83)	20.6	(2.26)	.955	(.04)
y_{12}: Index of age composition	9.37	(1.80)	26.3	(2.12)				
Exogenous variables								
z_2: Adolescent schooling	.318	(3.41)	.396	(2.85)	.273	(2.99)	.360	(2.65)
z_3: Index of demand for female labor services	.0155	(.26)	.0988	(.73)	.0217	(.35)	.0468	(.36)
z_7: Unemployment	.281	(1.04)	.0614	(.20)	.347	(1.28)	.191	(.63)
Constant	-108.	(2.12)	271.	(2.08)	7.6	(1.89)	3.	(.01)

Note:
- Variable not included in regression.

Table B-8

EQUATION 5: Y_5, PERCENT OF WOMEN (14+) CURRENTLY CONSENSUALLY MARRIED, 1960 DATA
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.369				.286			
F			4.15				3.93	
Endogenous variables								
y_3: Income (100 $)	-.304	(.98)	-.570	(.92)	-.846	(3.18)	-1.34	(3.26)
y_7: Sex ratio	8.31	(2.07)	22.7	(2.19)	7.91	(1.87)	22.3	(2.04)
y_{12}: Index of age composition	7.47	(2.98)	-8.11	(1.60)				
Exogenous variables								
z_2: Adolescent schooling	.0891	(1.30)	.0755	(.93)	.135	(1.92)	.13	(1.70)
z_3: Index of demand for female labor services	.0564	(1.37)	.113	(1.77)	.051	(1.17)	.13	(2.00)
z_7: Unemployment	.249	(.99)	-.044	(.13)	.307	(1.16)	-.04	(.14)
Constant	62.1	(2.66)	54.7	(1.15)	-6.07	(1.26)	20.1	(1.92)

Note:
– Variable not included in regression.

Appendix C

REGRESSION RESULTS FOR 1950 AND 1960 COMBINED EXCLUDING SAN JUAN, EQUATIONS 2 THROUGH 5

Table C-1

EQUATION 2: Y_2, FEMALE LABOR FORCE PARTICIPATION RATE, 1950 AND 1960 DATA COMBINED, SAN JUAN EXCLUDED
(t ratios in parentheses)

Type of Estimate	OLS		IV		OLS		IV	
	(1)		(2)		(3)		(4)	
R^2	.526				.523			
F			2.57				4.34	
Endogenous variables								
y_3: Income (100 $)	.137	(.29)	-5.35	(1.25)	.143	(.31)	-3.54	(1.18)
y_4: Legal marriage	.0860	(.28)	1.55	(.75)				
y_5: Consensual marriage	-.0422	(.15)	.371	(.24)				
y_8: Children under 5 per woman 15-49	-11.0	(2.14)	-46.8	(1.11)	-11.5	(2.43)	-13.8	(.64)
y_9: Index of age composition	-1.89	(1.02)	9.05	(.61)	-1.29	(.83)	12.0	(1.13)
Exogenous variables								
z_2: Adolescent schooling	-.288	(3.46)	.229	(.56)	-.314	(3.99)	-.155	(.77)
z_3: Index of demand for female labor services	.277	(5.62)	.486	(2.26)	.271	(6.12)	.433	(2.81)
z_7: Unemployment	-1.42	(5.32)	-1.31	(1.68)	-1.47	(5.66)	-1.85	(4.24)
Dummy (1960 = 1)	-1.96	(.74)	11.4	(.55)	-.623	(.28)	19.1	(1.29)
Constant	59.0	(1.81)	222.	(.76)	52.2	(1.64)	-219.	(1.02)

Note:
– Variable not included in regression.

Table C-2

EQUATION 3: Y_3, INCOME PER PERSON OF LABOR FORCE AGE, 1950 AND 1960 DATA COMBINED, SAN JUAN EXCLUDED

Type of Estimate	OLS		IV		OLS			
	(1)		(2)		(3)		(4)	
R^2	.873				.872			
F								
Endogenous variables								
y_2: Female labor force participation rate	-.0163	(1.03)	-.128	(1.21)	-	-	-	-
y_4: Legal marriage	.141	(3.32)	-.212	(.80)	.141	(3.32)	.376	(1.42)
y_5: Consensual marriage	.115	(2.50)	-.341	(1.23)	.118	(2.57)	.462	(1.54)
y_8: Children under 5 per woman 15-49	-.968	(1.04)	-3.54	(.64)	-.820	(.89)	4.47	(.70)
y_{10}: Sex ratio	4.88	(4.95)	9.90	(1.34)	5.03	(5.16)	6.1	(2.62)
Exogenous variables								
z_4: Adult education	.561	(4.55)	.298	(1.23)	.581	(4.77)	.361	(1.32)
z_7: Unemployment	-.0835	(1.61)	-.265	(1.57)	-.0570	(1.26)	-.0896	(.91)
z_8: Agriculture	-.0994	(10.2)	-.126	(2.44)	-.0935	(11.8)	-.0701	(2.58)
Dummy (1960 = 1)	1.18	(5.20)	-.286	(.31)	1.28	(6.27)	.661	(1.13)
Constant	-3.59	(1.50)	2067.	(1.86)	-4.67	(2.17)	8.	(1.44)

Note:
- Variable not included in regression.

Table C-3

EQUATION 4: Y_4, PERCENT OF WOMEN (14+) CURRENTLY LEGALLY MARRIED, 1950 AND 1960 DATA COMBINED, SAN JUAN EXCLUDED
(t ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.484				.47[illegible]			
F			.44				11.10	
Endogenous variables								
y_3: Income (100 $)	1.07	(3.06)	2.75	(2.67)	1.32	(4.22)	1.69	(2.38)
y_7: Sex ratio	3.45	(.75)	-3.69	(.21)	-4.37	(.96)	9.83	(.70)
y_{11}: Index of age composition	.872	(1.65)	-2.44	(1.54)	-			
Exogenous variables								
z_2: Adolescent schooling	-.251	(4.36)	-.315	(3.92)	-.262	(4.57)	-.314	(4.24)
z_3: Index of demand for female labor services	-.116	(3.31)	-.136	(1.69)	-.108	(3.08)	-.120	(1.63)
z_7: Unemployment	-.329	(1.74)	-.263	(1.10)	-.339	(1.78)	-.262	(1.19)
Dummy (1960 = 1)	4.42	(2.84)	2.88	(.59)	2.86	(2.30)	3.14	(1.14)
Constant	20.3	(.75)	186.	(2.11)	64.4	(13.0)	51.8	(3.73)

Note:
- Variable not included in re

Table C-4

EQUATION 5: Y_5, PERCENT OF WOMEN (14+) CURRENTLY CONSENSUALLY MARRIED, 1950 AND 1960 DATA COMBINED, SAN JUAN EXCLUDED
(t ratios in parentheses)

Type of Estimate	OLS (1)		IV (2)		OLS (3)		IV (4)	
R^2	.516				.513			
F			0.78				9.89	
Endogenous variables								
y_3: Income (100 $)	-.560	(1.49)	[illegible].92	(2.00)	-.729	(2.35)	-1.85	(2.63)
y_7: Sex ratio	11.0	(2.43)	[illegible].0	(.83)	.112	(2.47)	10.3	(.74)
y_{12}: Index of age composition	-2.11	(.81)	[illegible].0	(1.37)				
Exogenous variables								
z_2: Adolescent schooling	.226	(3.89)	.373	(2.61)	.235	(4.14)	.240	(3.28)
z_3: Index of demand for female labor services	.0245	(.71)	.196	(1.65)	.0256	(.74)	.117	(1.62)
z_7: Unemployment	.329	(1.74)	.0361	(.11)	.327	(1.73)	.156	(.72)
Dummy (1960 = 1)	-6.80	(1.76)	[illegible].8	(1.05)	-5.79	(4.71)	-2.98	(1.09)
Constant	12.9	(.50)	-294.	(1.40)	-7.46	(1.52)	-8.32	(.61)

Note:
- Variable not included in regression.

Appendix D

REGRESSION RESULTS FOR EQUATION 1 WITH INFANT DEATH RATE AS THE MEASURE OF MORTALITY REGIME

Table D-1

EQUATION 1: Y_1, CRUDE BIRTH RATES WITH INFANT DEATH RATES, z_9 AS MEASURE OF MORTALITY REGIME

	Variable Levels		Deviations from Municipio Means		Transformed Variables	
Type of Estimate	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
R^2	.239		.551		.532	
F		16.88		18.70		16.80
Endogenous variables						
y_2: Female labor force participation rate	-.0215 (.61)	.161 (1.02)	.0525 (1.76)	-.376 (1.95)	-.0269 (.64)	-.241 (1.95)
y_3: Income (100 $)	-.731 (4.53)	3.26 (1.65)	-.742 (4.07)	2.61 (2.27)	-.564 (3.33)	1.15 (.69)
y_4: Legal marriage	.348 (3.74)	-3.01 (3.28)	.993 (7.84)	2.36 (2.63)	.787 (7.16)	2.84 (2.56)
y_5: Consensual marriage	.530 (5.97)	-.766 (.99)	1.00 (10.6)	3.18 (6.09)	.932 (9.93)	3.00 (6.58)
y_6: Index of female age composition	.185 (3.19)	-1.91 (2.29)	.142 (2.49)	-.650 (2.72)	.191 (3.47)	-.109 (.28)
Exogenous variables						
z_2: Adolescent schooling	.0208 (.59)	-1.08 (3.00)	-.0557 (1.29)	-.272 (2.46)	-.0269 (.64)	-.0767 (.67)
z_4: Adult education	.899 (3.23)	-.533 (.49)	-.641 (2.47)	3.30 (4.15)	-.523 (2.02)	-2.05 (2.01)
$z_9(-3)$: Infant death rates lagged three years	.00989 (.90)	-.00612 (.21)	.00931 (1.61)	.00160 (.18)	.0106 (1.81)	.00476 (.55)
$z_9(-4)$: Infant death rates lagged four years	.00193 (.18)	-.00774 (.28)	.0117 (2.08)	-.00367 (.40)	.0129 (2.26)	.00074 (.08)
z_6: Unpaid family workers	.688 (6.59)	1.82 (4.07)	-.369 (4.27)	-.651 (4.37)	-.304 (3.48)	-.609 (4.53)
Constan	5.21 (.79)	273. (3.72)	a	a	-3.95 (1.85)	-35.2 (1.88)

Table D-1 (continued)

Type of Estimate	Variable Levels		Deviations from Municipio Means		Transformed Variables	
	OLS	IV	OLS	IV	OLS	IV
	(7)	(8)	(9)	(10)	(11)	(12)
R^2	.229		.548		.525	
F						
Endogenous variables						
y_2: Female labor force participation rate	.00468 (.14)	.127 (1.36)	.0620 (2.09)	-.353 (2.19)	-.0902 (3.12)	-.226 (1.90)
y_3: Income (100 $)	-.604 (3.84)	-.877 (1.82)	-.704 (3.86)	.936 (1.15)	-.475 (2.82)	.822 (1.09)
y_4: Legal marriage	.321 (3.44)	-1.91 (4.11)	.993 (7.82)	2.53 (3.39)	.750 (6.80)	2.99 (4.71)
y_5: Consensual marriage	.511 (5.73)	.979 (2.14)	1.06 (11.4)	2.49 (6.53)	.988 (10.7)	2.98 (6.58)
y_6: Index of female age composition						
Exogenous variables						
z_2: Adolescent schooling	.00114 (.03)	-.313 (4.01)	-.0749 (1.75)	-.191 (2.15)	-.0512 (1.23)	-.0410 (.58)
z_4: Adult education	1.02 (3.70)	.542 (.94)	-.710 (2.74)	-2.18 (3.83)	-.580 (2.23)	-1.83 (2.96)
$z_9(-3)$: Infant death rates lagged three years	.0139 (1.26)	-.0253 (1.53)	.00893 (1.54)	.00599 (.82)	.0108 (1.82)	.00550 (.70)
$z_9(-4)$: Infant death rates lagged four years	.00469 (.44)	-.0214 (1.35)	.0107 (1.91)	.00454 (.63)	.0122 (2.12)	.00174 (.23)
z_6: Unpaid family workers	.730 (7.00)	.978 (6.43)	-.386 (4.47)	-.560 (4.62)	-.321 (3.65)	-.607 (4.65)
Constant	11.4 (1.80)	151. (5.03)	a	a	-1.36 (.68)	-38.9 (3.97)
			$\hat{\rho}$ = .890	$\hat{\rho}$ = .921		

Notes:

[a]Intercept suppressed.

–Variable not included.

Appendix E

Table E-1

MEANS AND STANDARD DEVIATIONS OF DERIVED VARIABLES USED IN REGRESSION ANALYSI

Derived Variables	1950 Mean	1950 Standard Deviations	1960 Mean	1960 Standar Deviati
Endogenous				
y_1	37.7	4.49	32.9	5.57
y_2	18.6	8.48	16.1	4.90
y_3	4.74	1.46	6.87	2.32
y_4	44.4	5.30	47.7	3.94
y_5	14.5	5.37	8.67	3.55
y_6	32.8	2.44	30.9	4.54
y_7	.996	.0945	.927	.122
y_8	.804	.0899	.737	.135
y_9	20.4	.273	19.5	.443
y_{10}	1.02	.109	.946	.141
y_{11}	51.0	.800	49.7	.975
y_{12}	9.71	.135	9.36	.210
Exogenous				
z_1	67.8	6.18	82.7	5.77
z_2	45.6	8.37	52.9	7.27
z_3	78.5	21.6	81.2	18.5
z_4	3.34	.743	3.96	.967
$z_5(-3)$	11.6	2.34	6.76	1.26
$z_5(-4)$	13.2	2.68	6.98	1.31
z_6	3.72	2.65	1.25	1.21
z_7	4.03	2.33	5.57	1.66
z_8	49.8	15.2	36.3	14.9

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SUMMARY OF DERIVED VARIABLES

Symbol	Brief Description
Endogenous Variable	
y_1	Crude birth rate
y_2	Female labor force participation rate
y_3	Personal income (100 $)
y_4	Legal marriage rate
y_5	Consensual marriage rate
y_6	Age index of birth rates
y_7	Sex ratio (M/F) age 15 - 34
y_8	Children (0-4) per woman 15 - 49
y_9	Age index of female participation rates
y_{10}	Adult sex ratio (M/F) 15-64
y_{11}	Age index legal marriage rates
y_{12}	Age index consensual marriage rate
$y_{13} - y_{24}$	Migration variables (see Table 2)
Exogenous variables	
z_1	Schooling rate, age 7-13
z_2	Schooling rate, age 14-17
z_3	Index of demand for female labor services
z_4	Adult schooling (years)
z_5	Crude death rate
z_6	Unpaid family workers
z_7	Unemployment rate
z_8	Agriculture
z_9	Infant death rate