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THE EFFECT OF PIPED WATER ON EARLY CHILDHOOD MORTALITY IN URBAN BRAZIL, 1970 TO 1976

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This paper analyzes the impact of access to piped water on trends and income class differentials in child mortality in Brazil. Starting with a review of what available data indicate about general trends in infant mortality since 1950, it then concentrates on changes in urban areas between 1970 and 1976, using census data for the first and survey data for the second date. For mothers aged 20–29, path analytic regression techniques are used to test a recursive model linking the supply and demand for piped water to selected household and community level variables, and to examine their joint effects on child mortality. The model serves two related objectives: first, to examine the role of water supply in child mortality differentials at two points in time, 1970 and 1976; and second, to determine how much increased access to water between those dates contributed to (a) a decline in the level of child mortality and (b) a narrowing of income class differentials in mortality. Since income class differentials in mortality in Brazil and other developing countries are a growing concern (see Behm, 1980), and increased access to safe water is a public sector intervention that could reduce them, these questions have significant policy as well as theoretical interest.

The effectiveness of such interventions depends both on the capacity of implementing agencies to supply the appropriate types of services to low income families, and on those families' capacities to marshall their limited resources to utilize services to improve their children's chances of survival. The theoretical implication of this observation, as Selowsky (1981, 8–11) has pointed out, is that there is a demand as well as a supply side to the question. Demand is understood not in the sense that poor parents would choose between having more or fewer of their children survive, but rather that the success of their efforts to raise healthy children depends on the choices they make in allocating their limited resources.

The Brazilian experience is particularly interesting because two developments relevant to child survival are present. First, during the early 1970s that country embarked on an intensive effort to improve urban environmental conditions with a known link to mortality: water supply and sanitation. PLANASA, the national water and sanitation program, provided public subsidies designed to supply 80 percent of urban households with piped water and sewer connections by 1980 (Almeida, 1977). Second, Brazil made substantial progress in raising the level of educational attainment during the 1960s, including that of women; these women became the mothers of the 1970s and maternal education is an important determinant of the survival chances of infants and young children (Behm, 1980; Caldwell, 1980). These developments raise the question of whether interclass differentials in infant and child mortality were affected by such measures, and of how their effectiveness in reducing mortality may have been conditioned by continued income inequality.

The analysis is based on subsamples extracted from both the one-percent sample of Brazil's 1970 population census (Brazil, Fundação IBGE, 1979) and the 1976 PNAD survey (Brazil, Fundação IBGE, 1981b). The study populations consist of currently married (with husbands present) urban women between the ages of 20 and 29 who had had at least one live birth. In addition to data on these women's births, survival of their children, and their education, the subsamples include information on their husbands and on the households in which they resided. The choice of these samples is based on several considerations. The PLANASA program was an urban project; the study focuses on various types of urban water supply rather than rural-urban differences.[1] Access to water was observed at the time of the interview, while data on the survival of children relate to a period of years before the interview; limiting the analysis to women in their twenties reduces the time differential between the two variables, since this would increase with a mother's age, but it does not eliminate the problem.[2] Finally, living arrangements of mothers separated from their husbands were more varied and require more complex specifications in order not to distort the analysis. Because the subsamples are restricted, however, the results must be interpreted cautiously with reference to the Brazilian population as a whole.

TRENDS AND DIFFERENTIALS IN CHILD MORTALITY IN BRAZIL

Most of the available evidence on trends and differentials in child mortality in Brazil consists of indirect measures derived from survival ratios in the censuses and PNAD surveys.[3] Table 1 presents estimates of trends and regional differentials in infant mortality reported in the National Academy of Sciences' report on Brazilian fertility and mortality levels and trends, along with rates based on advanced tabulations of the 1980 census, which became available after the National Academy's estimates were prepared. (For the geographic composition of the PNAD regions, see the appendix figure.) At the national level, infant mortality declined from 130 infant deaths per 1000 live births in the years before the 1950 census to 98 in the 1970 census, a decrease of 25 percent. The decline was greater during the 1960s (18 percent) than the 1950s (8 percent). Regional differentials (using the percentage difference between infant mortality rates in the Northeast and São Paulo as a rough index of such differentials) increased from 46 percent in 1950, to 62 percent in 1960, and to 73 percent in 1970.

After 1970 the pace of the national decline slowed, with the rate falling by 12 percent between the late 1960s and late 1970s. The 1972, 1973, and 1976 PNAD surveys and the 1974–75 ENDEF study present conflicting stories about what may have been occurring during the decade. The PNAD surveys suggest an acceleration of mortality decline during the first half of the decade, with a slowdown after 1976. The ENDEF results suggest that very little decline occurred during the early 1970s, though caution is required in comparing 1970 census data with ENDEF because of the nature of that survey.[4] Comparison of PNAD results with preliminary 1980 census figures suggests either that the PNADs may have overstated the decline in the early 1970s or that conditions leading to accelerated decline in the first half of the decade may have changed for the worse in the second half.

Income Class Differentials in Mortality

Significant differentials are an important feature of Brazilian mortality patterns. They have been observed among regions and income classes, particularly for infants and children (Carvalho and Sawyer, 1978; Brazil, Fundação IBGE, 1981a). Infant mortality rates observed in a number of Brazil's larger cities increased from the late 1960s to the early 1970s, leading some observers to hypothesize that this reflected declining living standards among lower income

Table 1.—Estimated infant mortality per 1000 live births[a], Brazil and PNAD regions, 1950–1980

	Year of census or survey[b]							
Region	1950 Census	1960 Census	1970 Census	1972 PNAD	1973 PNAD	1974–1975 ENDEF	1976 PNAD	1980 Census
Brazil[c]	130	119	98	100	98	108	89	86
1. Rio de Janeiro	114	99	79	70	69	85	61	70
2. São Paulo	112	97	79	66	65	79	66	64
3. Southern states	100	90	75	68	66	74	69	76
4. Minas Gerais, Espirito Santo	118	107	91	79	80	93	74	79
5. Northeastern states	164	157	137	141	141	148	122	116
6. Brásilia	-	110	83	80	79	84	65	67
7. Amazon states	119	105	88	-	-	-	-	72

[a]Estimates derived by using model life table techniques and based on Coale and Demeny (1983) model South.

[b]The reference period for mortality rates derived through indirect estimation is generally two to four years prior to the date of the census or survey interview. For estimates of the time reference in the 1970 and 1976 data, see Table 2.

[c]PNAD data exclude rural areas of Amazon states.

Sources: NAS (1983), PNAD survey reports, and preliminary tabulations of the 1980 Census.

groups (Yunes and Ronachezel, 1974; Wood, 1977).[5] Carvalho and Wood (1978) found substantial differences in life expectancy in their tabulations of 1970 data by household income level, with a striking twenty-five year gap between the life expectancy of the lowest income class in the Northeastern region and the highest income class in the Southeast. Since raw data files for both the 1970 census and the 1976 PNAD survey were available, ratios were retabulated using household income decile categories derived by Lluch (1981). The decile categories were then clustered into three groups (deciles 1-3, 4-6, 7-10), following a suggestion by Carvalho and Wood that possible bias arising from interclass mobility in rates based on retrospective reporting can be reduced by grouping the population into fewer and broader categories.[6]

Trussell's variant of the Brass model life table method of calculating child mortality rates from questions on surviving children was used to derive probabilities of dying between birth and ages two and three (${}_2q_0$ and ${}_3q_0$) for children born to women grouped according to their household income decile in 1970 and 1976, as shown in Table 2.[7] Focusing on the average decrease in the probability of dying between birth and ages two and three, the table shows an overall decline of 26 percent from 1970 to 1976. This is greater than the 17 percent decline in the national level infant mortality rate reported in Table 1 and reflects the restriction in the study population to the group described above. Declines were greater for higher income groups in both rural and urban areas, but urban declines for all groups were greater than those in rural areas (the upper income groups in

Table 2.—Estimated probabilities of death by ages two (${}_2q_0$) and three (${}_3q_0$) by household income deciles, rural and urban areas, 1970 and 1976

Household income deciles	1970					1976					Average percent decline
	${}_2q_0$	(t)	${}_3q_0$	(t)	Percent of women	${}_2q_0$	(t)	${}_3q_0$	(t)	Percent of women	
Total	.133	(2.06)	.138	(3.78)	100.0	.099	(2.16)	.100	(3.83)	100.0	26
Urban											
1-3	.191	(2.13)	.217	(3.77)	10.8	.137	(2.53)	.161	(4.19)	12.2	27
4-6	.143	(2.19)	.144	(4.08)	16.5	.104	(2.11)	.108	(4.02)	18.6	26
7-10	.097	(1.93)	.103	(3.29)	35.8	.066	(1.86)	.061	(3.25)	37.7	37
Rural											
1-3	.154	(2.22)	.166	(4.13)	17.8	.138	(2.43)	.139	(4.32)	13.4	13
4-10	.131	(1.90)	.126	(3.39)	19.1	.106	(1.96)	.104	(3.41)	18.1	18

(t) is the approximate reference point in years prior to the interview.

Source: 1970 and 1976 sample files

rural areas were consolidated because of the limited number of cases in the 7-10 category). Timing may be involved in income class differentials, since women in higher income groups are generally older when they begin childbearing. The reference period for mortality rates recorded for them is shorter and should therefore capture more decline than is recorded for women in lower income categories. Even if the timing effect could account for the apparent widening of interclass differentials, however, the data in Table 2 raise doubts about the effectiveness of efforts to reduce mortality differentials during the 1970s since they do not suggest that the differentials narrowed.

DETERMINANTS OF MORTALITY DIFFERENTIALS

Variables that influence infant and child mortality and that could explain differentials among income groups include those with direct as well as indirect effects on the survival chances of young children. Relevant direct causal variables include the nutritional level of both mothers and children, the quality of their environment, and the maternal and child health care to which they have access (Dyson, 1978; Antonovsky, 1980; Gwatkin et al., 1980; Loriaux, 1980). Background variables such as income and education may influence mortality directly or by affecting or conditioning directly causal variables.

Empirical investigations of the determinants of infant and child mortality have suggested that causal relationships are reinforcing rather than distinct and separable. Puffer and Serrano (1973) studied infant and child mortality patterns in several Latin American cities, including three in Brazil (Recife, Riberao Preto, and São Paulo). Diarrheal diseases were the leading causes of postneonatal deaths in all three cities, with the proportion being highest in the poorest of the three, Recife. The investigation revealed a strong synergism between poor nutrition and environmental factors, and concluded that children who were underweight at birth owing to poor maternal nutrition and who were poorly nourished after birth were also more susceptible to infections associated with unsafe water, poor sanitation, and inadequate housing.[8] Their study also showed that a high proportion of the households reporting infant deaths lacked adequate water sources, though they did not analyze systematically the relation between the two variables.

Access to internal piped water in the household is likely to be of most direct benefit in lowering child mortality by reducing exposure to water-borne diseases, particularly diarrheal disorders. Although evidence of the specific links between water supply and such disorders is limited, the literature suggests that increases in the amount of water used contribute to better hygiene and that elimination of bacteriological contamination reduces the risk of infection through intake. Schneider, Shiffman, and Faigenblum (1978) present clinical data from a Guatemalan study that document the loss of nutrients during diarrheal episodes, and a recent World Bank study reports that gastrointestinal diseases reduce absorption of nutrients by as much as 30 percent in acute cases (World Bank, 1980). An added dimension in the Brazilian situation relates to the very low average duration of breastfeeding (Anderson, 1981). Thus infants are exposed to the risk of infection from contaminated water and through poor hygiene in prepared food at a very early age.

This synergism extends to other socioeconomic variables, particularly maternal education (Behm, 1978; Arriaga, 1980; Cochran et al., 1980). Because the causal influences may be both direct and indirect, explanations of the way in which maternal education affects mortality differentials vary. In interpreting Nigerian data, Caldwell (1980, 13-14) suggests three links: (a) mothers with more education tend to be less fatalistic about illness and therefore prone to seek outside medical assistance for an ill child; (b) educated mothers are more likely to adopt improved child care practices such as boiling water used in the preparation of infant formulas; and (c) education may change intrafamily relationships, leading to a more "child-centered" orientation that would have a positive impact on children's health.

Schultz (1980), drawing on household production theory, hypothesized that better educated women earn more in the labor market and marry better educated men; consequently they have higher family incomes enabling them to purchase goods and services that improve child health. Education may also increase the effectiveness of women's nonmarket child care activities, although, as Schultz cautions, the fact that market work requires women to be absent from the home could have an offsetting negative impact on the quality of child care.

Brazilian data illustrate the strong link between maternal education and children's survival chances. Table 3 shows the proportion dying among children (all ages) of urban mothers by educational level of mother. Education categories range from women with no formal education to those with ten or more years of schooling as reported in the 1970 census and the 1976 PNAD survey. Current age of the mother is used as a control variable. There is a steady decline in the proportion of children dying as educational attainment increases. The most striking difference in mortality ratios, however, is the contrast between mothers with no formal education and other groups: the rates of child mortality among women with ten or more years of schooling range from a third to less than a tenth of those for mothers with no education. Rates in all education groups are generally lower in 1976 than in 1970. There is a substantial decline in the relative number of women with no formal education, and there are proportional increases in the two highest classes, reflecting the increased emphasis on education mentioned earlier. The changing educational distribution is consistent with other data from the Brazilian censuses which show that 41 percent of women aged 14–18 in 1970 (the 20–24 year olds in 1976) were enrolled in school at that date compared to an enrollment of only 21 percent among 14–18 year olds in 1960. The last column of Table 3 shows a decline in the proportion of children dying for mothers of all ages, a finding consistent with the improved survival probabilities shown in Table 2.

SPECIFICATION OF A CAUSAL MODEL

The census and PNAD provide information on the education, the earnings and the employment status of household members, and on a number of household characteristics, including access to piped water, thereby permitting a more detailed examination of at least some aspects of the link between increased access to piped water, recent social and economic changes in Brazil, and trends and differentials in child mortality. The measure of child mortality analyzed below, CM(i), where "i" refers to individual mothers, is based on the ratio of surviving children to births for the mothers in the study population. It will be discussed in detail in the next section. For each mother, a dummy variable, WAT(i), indicates whether the household in which she resided had access to piped water at the time of the interview.

An adequate causal model of the links between improved access to piped water and child mortality needs to take account of two sets of relationships. One set pertains to variables associated with the consumption of this particular service by urban households. The other is concerned with the impact that this consumption, along with other relevant variables, has on the survival chances of children. Empirical testing of such a model requires specification of structural

Table 3.—Proportion of children who have died by mother's age and years of schooling, urban areas, 1970 and 1976

Age	Years of Schooling None	1-4	5-9	10+	All groups
			1970		
20-24	.201	.106	.092	.034	.131
25-29	.192	.107	.083	.063	.127
30-34	.220	.127	.114	.048	.150
35-39	.219	.170	.127	.066	.176
40-44	.262	.168	.076	.055	.184
45-49	.301	.180	.127	.094	.224
Percent	26.8	47.3	19.0	6.9	100.0
			1976		
20-24	.193	.098	.053	.029	.094
25-29	.134	.093	.071	.029	.089
30-34	.211	.096	.085	.019	.110
35-39	.193	.110	.067	.029	.121
40-44	.222	.119	.076	.068	.143
45-49	.232	.145	.082	.052	.163
Percent	16.6	46.6	23.4	13.4	100.0

Source: Sample files

equations for both sets of relationships, as well as the links between them.

Consumption has both a supply and a demand dimension. As Selowsky (1981) has illustrated in his study of the consumption of services in Colombia, a household may or may not consume a service such as water, (a) because the supply network is not geographically accessible or (b) because, in spite of accessibility, that household chooses not to do so for reasons of price, income limitation, or as a matter of preference. The former is a supply constraint, the latter demand.[9] A specification that accounts for supply variables requires information on availability of water at a comparatively low level of aggregation. Selowsky (1981) worked with neighborhoods in his study of Colombia. The level of aggregation is much greater in the Brazilian data. The public use sample of the 1970 census permitted tabulations for 117 geographic areas, consisting of clusters of municipalities. A water supply variable, WS(j), where "j" refers to the geographic area of residence, was calculated as the proportion of households in each of the 117 areas reported to have internal piped water in 1970. Since the 117 areas were not identifiable in the PNAD survey, WS(j) was calculated for each of

Brazil's 26 states in 1976. Because of the level of aggregation, WS provides only a rough measure of differences in local water availability. A second variable, PWS(j), measuring the change in the proportion of households served in each state between 1970 and 1976, was added as a measure of the impact of the PLANASA program on the supply of water. Water supply was assumed to be exogenous at the level of individual households, with the added restriction that the effect of supply on mortality at the household level was indirect (via WAT).

Data limitations also hampered specification of demand variables. While the data files included information on a number of household level variables (e.g. income and education) that influence demand, no information was available at any level of aggregation on price (hookup and use fees) differentials. Further, those household level variables that determine the demand for water are also likely to have a causal relation with child mortality in their own right, and one (income) is endogenous at the household level. Since the study population consists of mothers, with husbands present, in single family households, and only a small proportion (.08) of those women were working, the husband's earnings, HY(i), is used as a measure of the endogenous family income variable. Education of husbands, HED(i), and mothers, MED(i), are assumed to be exogenous at the household level. Identification of the state of residence in the two data files also made it possible to add state level information from other sources: a variable measuring per capita value added taxes, TAX(j), in each state serves as a proxy for regional differences in labor market conditions.

The causal relations between mortality differentials, access to piped water, and the household and community level determinants of those variables that this paper seeks to test can be summarized in the following set of structural equations:

$$[1]\ HY(i) = f[HED(i), TAX(j)]$$

$$[2]\ WAT(i) = w[MED(i), HED(i), HY(i), WS(j)]$$

$$[3]\ CM(i) = c[MED(i), HED(i), HY(i), WAT(i), TAX(j)]$$

Equation [1] specifies the first of the three endogenous variables in the model, husband's earnings (HY), in terms of his education, HED, and the value added tax per capita in the state in which he resides, TAX. While HED reflects the influence of individual characteristics on earnings, the value-added tax is an argument in the earnings function because it reflects regional differences in the productivity of manufacturing industry in Brazil and their impact on earnings. The second endogenous variable, WAT, is represented in equation [2] as a function of household variables related to demand for water (HED, MED, HY) as well as the community level water supply variable, WS. The variable measuring increases in water supply between 1970 and 1976, PWS(j) was added to the 1976 equation:

$$[2a]\ WAT(i) = w[HED(i), MED(i), HY(i), WS(j), PWS(j)]$$

Finally, equation [3] represents child mortality, CM, as a function of exogenous household variables (HED and MED), endogenous household variables (HY and WAT), and the exogenous community level variable, TAX. The latter is included in [3] because there is good reason to expect that the regional differences in economic conditions that TAX measures not only affect CM indirectly through HY and WAT, via equations [1] and [2], but also directly through its effects on other aspects of the standard of living in the community.

Since the arguments on the right hand side of equation [3] also appear in equations [1] and [2], ordinary least squares regression will not yield estimates of structural coefficients without further restrictions (Marsden, 1981, 118). Two such restrictions are invoked, one on the basis of reasoning about the nature of the causal relations expressed in the system, and the other based on an empirical check on the error terms in the estimating equations. The first restriction is that the model is fully recursive in terms of the relations among endogenous variables: HY is causally prior to WAT, and both are causally prior to CM. The alternative would be to argue that there is a simultaneous relationship among the three variables, which does not appear to be warranted in this case. The second restriction is that the error terms of the three estimating equations be uncorrelated with each other and with the right hand side variables. Empirical tests showing that this restriction is justified will be reported below.

With these restrictions, ordinary least squares (OLS) estimates of the reduced form coefficients of the exogenous variables can be used to estimate their total effect on CM, and path analytical techniques can be utilized to decompose their total effect into direct effects on CM and their indirect effects via HY and WAT, as well as the total and direct effects of the latter (Alwin and Hauser, 1975).[10]

Measurement of the Dependent Variable

Several methodological problems arise in the use of ratios of children surviving to children ever born in analyses of differences in mortality at the level of individual mothers. Indirect measurement techniques that are used to translate survival ratios into mortality rates for groups of women cannot be used to assign a mortality rate to an individual woman. Differences in survival ratios among individual mothers reflect not only the particular mortality regime to which their children are subject, but also differences in these children's exposure to the risk of dying, which is related to the number and timing of the births that the woman has had. This, in turn, is related to her age and to the duration of her marriage or other type of union. Finally, the relation of survivors to births is binomially rather than normally distributed, making OLS an inappropriate estimating technique unless the relation can somehow be normalized.

Trussell and Preston (1982) have suggested a method for standardizing survival ratios in individual level data that controls for duration of risk and normalizes the distribution of these ratios. Their method employs a model schedule of survival ratios to determine the expected number of infant deaths for women in different marital duration or age categories. Observed deaths are then divided by the number of deaths expected given the woman's age or marital duration and the appropriate model schedule, which yields an index that relates the individual's experience to the standard.[11] The authors have demonstrated that this index can be used in multiple regression analysis of individual level data on child survivorship. Nonetheless, tradeoffs are involved in using OLS rather than estimating techniques that take account of the fact that the distribution of the index is bounded by zero.[12]

Table 4 illustrates how the index of observed/expected deaths can be used to measure differences in the child mortality ratio, CM, by educational attainment and income for various types of water supply. In the table, CM is cross-tabulated by categories of mother's years of schooling, MED, as well as by the husband's income decile and the various types of water to which households in the sample had access, including internal piped water, the water variable used in the analysis. The first column for each of the two dates shows CM for women in households with internal piped water, followed by columns for women in households with water from various external sources.

The pattern of educational differences varies according to whether the water

Table 4.—Ratio of observed to expected child deaths, mothers aged 20–29, by years of schooling and income decile, by water supply source, urban areas, 1970 and 1976

	Water Supply				Percent of women in each category
		External			
	Internal piped	City system	Well/ spring	Other	
1970					
Women, aged 20-29	0.58	0.95	0.95	1.24	100.0
Education					
None	0.99	1.22	1.17	1.40	34.8
1-4	0.59	0.83	0.83	1.09	46.1
5-9	0.42	0.80	0.93	0.78	14.2
10+	0.30	0.84	*	0.78	4.9
Income decile					
1-3	0.84	1.01	1.04	1.30	(30)
4-6	0.63	0.80	0.99	1.21	(30)
7-10	0.53	1.07	0.81	1.10	(40)
N	3722	825	2086	3455	10088
Percent	36.9	8.2	20.7	34.2	100.0
1976					
Women, aged 20-29	0.37	0.77	0.85	1.22	100.0
Education					
None	0.73	1.43	1.10	1.54	12.3
1-4	0.46	0.70	0.83	1.18	43.1
5-9	0.31	0.43	0.69	0.69	28.4
10+	0.20	*	0.32	*	16.2
Income decile					
1-3	0.57	0.90	1.07	1.46	(30)
4-6	0.37	0.53	1.12	1.39	(30)
7-10	0.29	0.40	0.58	0.67	(40)
N	6312	1118	1325	987	9742
Percent	64.8	11.5	13.6	10.1	100.0

* Number of cases is less than 25.
Numbers in parentheses are approximate values.

Source: Sample files

source is internal or external. For women in households with internal water sources, the mortality index declines steadily with rising educational attainment. For external sources, there is a sharp difference between the "none" category and other groups, but it is generally less between the other groups. This suggests that while education may help to offset environmental disadvantages, its effect may be limited to such basic steps as boiling contaminated water. Beyond that, there appears to be greater complementarity between the effects of education and water. This suggests that the specification of equation [3] be modified to test for interactions between mother's education and water:

$$[3a]\ CM(i) = c[MED(i), HED(i), TAX(j), HY(i), WAT(i), D(i)]$$
$$\text{where } D(i) = MED(i) \cdot WAT(i).$$

Between 1970 and 1976, the share of women with no education declined from 35 percent to 12 percent of the sample population, which consists of women aged 20 to 29 rather than women aged 20 to 49 as reported in Table 3. The share of women with 10 and more years of schooling increased from 5 to 16 percent. The share of households with piped water increased sharply, from 37 percent to 65 percent. The pattern of differentials in child mortality by education and access to water also changed. Increased access to piped water appears to have offset a tendency for differentials in child mortality associated with educational attainment to increase. There were fewer mothers with no education in 1976 and, among those with internal piped water, the difference in their mortality index relative to other education classes declined. But among those with external water sources, the difference between mothers with no education and all other mothers increased.

In the second bank of Table 4, differences in the child mortality ratio by income deciles are generally greater in 1970 among households that have internal piped water than for those with outside water sources. In 1976, income differentials become clear among households with external sources. Underlying these changes is a shift in the composition of households relying on outside sources, with a substantial decline in the percentage of households with well/spring and other external sources. The relative improvement in the "city system" column for 1970 to 1976 may reflect improved water quality, since the PLANASA program improved the distribution system as well as access to it. These changes suggest that there may be interactions between income and water supply and/or a change in the nature of the relation between income and child mortality, in which case a test for income class/water supply interactions is warranted.

For all groups, the decline in child mortality is greatest for women in households with internal piped water, while there is practically no change for women in the "other" sources category and intermediate rates of decline in the groups with water from city taps and springs. Variability in child mortality among households without internal piped water is of interest because city taps and springs represent an intermediate (and less expensive) means of improving access to water. However, since the main effort in the PLANASA program was to increase access to internal piped water, and since most of the increased water access during the period under study occurred in that category, the analysis focuses on the contribution of that type of water access to differentials and changes in CM between 1970 and 1976.

ANALYSIS OF DIFFERENCES IN CM IN 1970 AND 1976

This section presents OLS estimates of the parameters in the causal model of determinants of child mortality shown in equations [1] to [3] using a path analytic approach. Given the recursive structure of the causal model, the analysis begins by estimating the reduced forms (including exogenous variables only) of equations for the three endogenous variables, and then proceeds to estimation of the structural equations in steps in which each of the causally prior endogenous variables is added to the regression equation of the final dependent variable, CM. The total, direct, and indirect effects of each variable on CM are then measured using standardized regression coefficients and conventional decomposition techniques (see Alwin and Hauser, 1975). A check on correlations among the error terms provides a test of the validity of the recursive specification.

Table 5 states the units in which the variables used in the regressions are measured, and provides sample means and standard deviations. Following the suggestion of Trussell and Preston (1982), observations are weighted by the number of children ever born in order to make the results representative of births rather than women. There is generally greater variability (as measured by the ratio of standard deviations to means) in the 1970 sample than in 1976, suggesting either increasing homogeneity in the population with respect to these variables between 1970 and 1976, or that the 1976 sample may be underrepresenting socioeconomic groups at the extremes of the range. An appendix table presents zero order correlation coefficients between each variable for both dates. The high correlation between MED and HED (0.6 to 0.7) reflects the tendency for the educated to seek mates with similar levels of education. Preliminary analyses of these variables indicated that their impact was non-linear, so that logarithmic transformations were used.

Regression results are presented in Table 6, following the sequence just described: columns (1), (2), and (4) are the reduced forms of equations [1], [2], and [3] of the causal model, while columns (3) and (6) are OLS estimates of structural equations [2] and [3] (there are no causally prior endogenous variables in

Table 5.—Weighted means and standard deviations, dependent and independent variables used in analysis of child mortality experience of urban mothers aged 20–29, 1970 and 1976

Variable	Description	1970 Mean (1)	1970 St.dev. (2)	1976 Mean (3)	1976 St.dev. (4)
Endogenous:					
CM	Standardized index	1.264	2.263	0.852	1.639
WAT	Dummy variable (=1 when household had piped water)	0.306	0.523	0.544	0.499
HY	Log of 1970 cruzeiros	5.104	1.300	6.091	0.885
Exogenous:					
MED	Log of years of school completed by mother (+1)[a]	0.842	0.920	1.356	0.780
HED	Log of years of school completed by father (+1)[a]	0.931	0.937	1.416	0.785
TAX	Log of 1970 cruzeiros	4.583	0.991	4.849	0.844
WS	Ratio of households in region with piped water to total households in region	0.419	0.240	0.467	0.191
PWS	Change in WS between 1970 and 1976	---	---	0.144	0.057

[a]To avoid taking the logarithm of zero, one was added to MED and HED in the log transformation of these variables.

Source: See text.

Table 6.—Husband's income (HY), piped water (WAT), and child mortality (CM), regressed on selected independent variables, 1970 and 1976[a]

Independent variable	HY (1)	WAT (2)	WAT (3)	CM (4)	CM (5)	CM (6)
			1970			
Constant	3.318	-.070	-.184	2.288	2.542	2.493
MED		0.088* (13.4)	0.081* (12.1)	-0.270* (8.4)	-0.252* (7.8)	-0.231* (7.1)
HED	0.579* (46.9)	0.095* (14.1)	0.082* (12.0)	-0.184* (5.7)	-0.156* (4.7)	-0.134* (4.0)
TAX	0.272* (23.3)	-0.013 (1.6)	-0.018 (2.1)	-0.104* (2.6)	-0.093 (2.3)	-0.098* (2.4)
WS		0.653* (18.1)	0.637* (17.7)	-0.350 (2.0)	-0.314 (1.8)	-0.145 (0.8)
HY			0.032* (7.8)		-0.071* (3.6)	-0.063* (3.1)
WAT						-0.264* (5.6)
R^2	.277	.237	.241	.046	.047	.050
F^b	1977.9	804.1	659.0	124.9	102.6	90.9
			1976			
Constant	4.121	-0.131	-0.694	2.493	2.626	2.424
MED		0.127* (18.5)	0.082* (12.0)	-0.234* (9.2)	-0.223* (8.5)	-0.200* (7.5)
HED	0.588* (63.6)	0.158* (23.0)	0.092* (13.0)	-0.261* (10.2)	-0.245* (8.9)	-0.219* (7.9)
TAX	0.234* (27.1)	-0.061* (3.9)	-0.015* (7.0)	-0.131 (2.3)	-0.120 (2.1)	-0.151* (2.6)
WS		0.962* (13.6)	1.010* (14.8)	-0.411 (1.6)	-0.263 (1.6)	-0.129 (0.5)
PWS		0.865* (10.3)	0.727* (8.9)	-0.890* (2.8)	-0.857* (2.7)	-0.646 (2.1)
HY			0.153* (24.6)		-0.036 (1.5)	-0.008 (0.3)
WAT						-0.290* (7.5)
R^2	.375	.271	.314	.067	.067	.073
F^b	2920.3	725.6	743.4	140.0	117.1	109.0

t-statistic in parentheses
* significant at .01 level

[a]Individual and household level data refer to mothers aged 20-29 resident in urban areas at time of inquiry.

[b]Degrees of freedom fo: .he F test are (NV, 9767 - NV - 1), where NV = the number of independent variables.

Source: See text and Table 5

equation [1] for husband's earnings). Column (5) includes the endogenous variable, HY, which is causally prior to WAT but omits WAT itself as a step in estimating indirect effects.

To avoid repetition, the results in Table 6 will not be described in detail since most of the conclusions that can be drawn from them are revealed when effects are decomposed in Table 7. It is worth noting that the addition of HY to the WAT regression had a greater effect in 1976 than in 1970, while the reverse occurred with the addition of HY to the CM equation. In fact, the HY coefficient was not statistically significant in 1976, suggesting a change in nature of the causal link between HY, WAT and CM. Another observation is that the coefficient for WS in the CM equations was not significant for either date, which is consistent with the point made earlier than the supply effect may be more indirect than direct. Finally, an analysis of error terms revealed no major problems with the recursive specification. Details are reported in a footnote.[13]

PATH ANALYSIS: TOTAL, DIRECT, AND INDIRECT EFFECTS

Decomposition of effects using standardized regression coefficients is useful first in revealing both the level and, since we have estimates for two points in time, changes in the relative contributions of variables to differences in child mortality, and second in providing a measure of the extent to which the effects of exogenous variables have been mediated by intervening endogenous variables. This is particularly appropriate for a study, which, in addition to seeking to determine the effect of access to piped water on child mortality, also tries to clarify the roles that both supply and demand variables affecting that access play in the relationship between water and child mortality.

Table 7 presents the decomposition of the effects on child mortality of variables in the regressions in Table 6. Total effects are shown in the first column. For the exogenous variables, total effects are the standardized coefficients from the reduced form regressions (column 4 in Table 6). For the endogenous variables, they are the coefficients from the first equation in which the variable enters (column 5 in Table 6 for HY and column 6 of Table 6 for WAT). In the case of WAT, the total effect equals the direct effect since that variable is the last in terms of causal ordering in the model. Indirect effects are shown in columns 2 (for HY) and 3 (for WAT), and were derived using the technique described by Alwin and Hauser (1975).

What does Table 7 tell us about: (*a*) the comparative importance of exogenous household and community level variables on differences in child mortality; (*b*) the extent to which their effect is direct, or mediated through endogenous household variables, WAT in particular; (*c*) the degree to which the mediation of WAT is influenced more strongly by household or community variables; and (*d*) changes, if any, in these relationships between 1970 and 1976? In considering (*c*) it is important to recall the dual roles assigned to the exogenous variables in the causal model that is being analyzed. They have a direct causal relation to child mortality, and indirect effects via HY and WAT, which have been specified as supply side effects on WAT for the community level variables and demand side effects for the household variables.

Mother's and husband's education, which are exogenous household level variables, have the largest total effect on child mortality in both 1970 and 1976. The total effect for MED is greater than for HED in 1970, but HED increases to slightly more than MED in 1976. Most of HED's increase is at the expense of the HY variable, which was statistically significant in the 1970 regressions (but not in 1976 (Table 6, column 6). The largest part of MED's effect is direct (the ratio of the direct effect to the total effect is

Table 7.—Total, direct and indirect effects of variables in equation [3] on child mortality (CM), 1970 and 1976

Independent variable	Total effect (1)	Indirect effect HY (2)	Indirect effect WAT (3)	Direct effect (4)
		1970		
MED	-.1099	-.0073	-.0087	-.0939
HED	-.0762	-.0117	-.0090	-.0555
TAX	-.0457	-.0049	+.0021	-.0429
WS	-.0371	-.0038	-.0178	-.0155
HY	-.0408	*	-.0048	-.0360
WAT	-.0612	*	*	-.0612
		1976		
MED	-.1113	-.0050	-.0114	-.0949
HED	-.1250	-.0073	-.0130	-.1047
TAX	-.0675	-.0055	+.0158	-.0778
WS	-.0479	+.0013	-.0342	-.0150
PWS	-.0311	-.0012	-.0073	-.0226
HY	-.0195	*	-.0150	-.0045
WAT	-.0882	*	*	-.0882

*No indirect effect

Source: See text and Table 5

.85 at both dates). The ratio of direct to total effect for HED is somewhat lower (.73) in 1970, but increases in 1976 (to .84), again reflecting the insignificance of the HY variable. The exogenous community level variables, TAX and WS, have a lower but still significant total effect on child mortality. They differ in the way in which their effects are mediated, since a larger share of the effect of TAX is direct, while less than half of the effect of WS is direct. In fact, the indirect effect of TAX via WAT is positive in sign, indicating that once demand is controlled there is a lower level of individual household access to piped water in areas with higher taxes per capita. Possible adverse effects of this on CM appear to be offset by *other* unspecified community level influences reflected in the negative direct effect of TAX. In the case of WS, a large share of its effect is mediated by WAT both in 1970 and in 1976, which is to be expected since it was introduced to measure the supply side impact of water availability at the community level on access at the household level. When community level *increase* in water supply (PWS) is introduced in 1976, a higher proportion of its effect is direct.

The total effects of the endogenous household level variables, HY and WAT, differ from those of the other two sets of variables. The total effect of WAT is greater than either of the two community level exogenous variables, but less than MED and HED. The effect of WAT increases from 1970 to 1976. HY's total effect is about the same as the community level variables in 1970 and slips well below them (and is, in fact, insignificant) in 1976.

In sum, the main changes from 1970 to 1976 are as follows: (*a*) the total effect of HED increases, as does the direct component of its effect; at the same time, the total effect of HY drops, while the indirect component of its effect (via WAT) increases (*b*) the total effect of water supply variables, WS in 1970 and WS plus PWS in 1976, increases, with most of the effect of that increase being mediated by WAT, the household level water variable. This increase in the indirect effect of exogenous supply via WAT suggests (*c*) that supply played a greater role in 1976 than in 1970, which is what one would expect the outcome of a successful public sector intervention to be. Interestingly, the total effect of the other community level exogenous variable, TAX, also increases, but none of that increase was mediated by the endogenous household level variables. As noted above, the indirect effect of TAX via WAT worked in the opposite direction, more so in 1976 than in 1970, which, given its interpretation as an indication that households in high TAX areas had lower access to WAT, also lends support to the conclusion that the importance of water supply at the community level increased from 1970 to 1976.

TEST FOR INTERACTION

An additional issue raised in relation to the causal model specified in equations [1]–[3] is interaction: the extent to which the effect of access to piped water on child mortality may have been greater or less for women within different education classes or in households with higher or lower income levels. The question of interactions is of particular interest because it would be useful to know whether a public sector intervention such as increased water supply has a redistributive effect on a sensitive measure of welfare, thereby compensating for some of the adverse effects of inequality in the distribution of human resources and income on the demand side.

As suggested earlier, equation [3] can be modified to test for interactions by adding interaction terms. Two such terms have been tested in both the 1970 and 1976 data, one for WAT and MED and the other for WAT and HY. If interaction is operating in the compensatory manner described above, then one would expect that differences in CM by MED and HY would be less in households with access to piped water than those without it, which would show up in the regressions in the form of flatter slopes for the MED and HY regression lines when WAT is present. Thus we would expect positive signs for both interaction terms.

Table 8 summarizes the tests for interaction. The first and fourth columns show the 1970 and 1976 regression results for equation [3] as reported in Table 6. The next two columns show the results when an interaction term is added, first for MED and then for HY. As expected, the signs of all four interaction terms are positive; however, only WAT/HY in 1976 is statistically significant (according to the "*t*" test, the probability that its coefficient is, in fact, zero is less than .02). This result suggests that while the slope of the overall HY regression line was zero in 1976, the slope for non-WAT households was negative, so that WAT did indeed contribute to a narrowing of income related differences in child mortality in 1976.

DECOMPOSITION OF CHANGES IN CM FROM 1970 TO 1976

A final question relates to the relative contribution of increased access to piped water to the 27 percent decline in child mortality between 1970 and 1976. Both the 1970 and 1976 regression equations can be evaluated using sample means for both dates to show the contribution of changes in those means to the change in the predicted value of child mortality between the two dates. This has been done in Table 9. To simplify the discussion of the role of water, regressions were rerun without the exogenous community supply measures, so that only WAT appears in the regression coefficients shown in the first column of the table. The second and third rows are the

Table 8.—Tests for interaction effects in equation [3a], dependent variable is child mortality (CM), 1970 and 1976[a]

Independent variable	1970 (interactions) None (1)	WAT/MED (2)	WAT/HY (3)	1976 (interactions) None (4)	WAT/MED (5)	WAT/HY (6)
CONSTANT	2.493	2.521	2.560	2.424	2.477	2.812
MED	-0.231*	-0.254*	-0.234*	-0.200*	-0.227*	-0.199*
	(7.1)	(6.8)	(7.2)	(7.5)	(6.8)	(7.5)
HED	-0.134*	-0.137*	-0.136*	-0.219*	-0.221*	-0.217*
	(4.0)	(4.1)	(4.1)	(7.9)	(8.0)	(7.8)
TAX	-0.098*	-0.099*	-0.100*	-0.151*	-0.149*	-0.146*
	(2.4)	(2.4)	(2.5)	(2.6)	(2.6)	(2.5)
WS	-0.145	-0.127	-0.123	-0.129	-0.128	-0.122
	(0.8)	(0.7)	(0.5)	(0.5)	(0.5)	(0.5)
PWS	-	-	-	-0.646	-0.644	-0.617
				(2.1)	(2.0)	(2.0)
HY	-0.063*	-0.065*	-0.082*	-0.008	-0.003	-0.067
	(3.1)	(3.3)	(3.4)	(0.3)	(0.1)	(1.6)
WAT	-0.264*	-0.338*	-.554*	-0.290*	-0.372*	-0.894*
	(5.6)	(4.6)	(2.8)	(7.5)	(5.3)	(3.3)
INT	-	0.069	0.054	-	0.063	0.102
		(1.3)	(1.5)		(1.4)	(2.2)
R^2	.050	.050	.050	.073	.073	.073
F^b	90.9	78.2	78.2	109.0	95.6	96.0

[a]t-statistic in parentheses
*significant at .01 level

[b]Degrees of freedom for the F test are as follows: 1970 data: (NV, 10356 - NV - 1); 1976 data: (NV, 9767 - NV - 1), where NV = the number of in dependent variables.

products of these coefficients and the sample means of variables in 1970 and 1976, respectively. Column 4 shows the differences between columns 2 and 3, which sum to the total change in CM. Column 5 expresses the ratio of the change contributed by each variable as a percent of the total change in CM.

Changes in the mean values of exogenous household characteristics account for the largest share of the total decline in child mortality. With 1970 regression coefficients, MED and HED account for 54 percent of the decline. HY, an endogenous household variable, accounts for another 18 percent. With 1976 coefficients, HY's contribution disappears (it shows up as a positive value, though it is statistically insignificant), and the combined contributions of MED and HED are 65 percent. Increased access to piped water plays a secondary but still significant role, contributing 18.7 percent to the decline using 1970 coefficients, and

22.3 percent with the 1976 results. While the 3.6 percentage point increase in the contribution of water to the decline in child mortality with 1976 coefficients does not signal a major structural change, the direction of the change lends further support to the view that the water program contributed to the amelioration of differences in child mortality that might have otherwise persisted owing to inequality in the distribution of income and human resources at the household level.

SUMMARY AND CONCLUSIONS

The objective of this paper has been to analyze the determinants of trends and differentials in early childhood mortality in urban Brazil during the 1970s. The mortality measure used in the study was the ratio, CM, of the actual proportion of children surviving for women in the study population to the proportion that would be expected to survive given a standard mortality schedule, those women's current age, and the ages at which they began to have children. The study population consisted of samples of currently married urban mothers extracted from Brazil's 1970 population census and its 1976 national sample survey (PNAD). The sample was restricted to women in their twenties in order to achieve a closer approximation between the timing of mortality estimates based on surviving children (which, on average, refer to a period two to four years prior to the date of interviews) and data on other characteristics that relate to the time of the interview. While most women in their twenties would have completed their education at an earlier date, their access to water could be very recent. This would lead to understatement of the true

Table 9.—Evaluation of OLS estimate of equation [3] to decompose change in child mortality (CM) from 1970 to 1976

Independent variable	Coefficient (1)	Coefficient x1970 means (2)	Coefficient x1976 means (3)	Change (3)-(2) (4)	Percent (4)/Total (5)
			1970		
Constant	2.5654	2.5654	2.5654	-	-
MED	-.2316	-.1949	-.3140	-.1191	34.4
HED	-.1377	-.1282	-.1949	-.0667	19.3
TAX	-.1249	-.5724	-.6054	-.0332	9.6
HY	-.0633	-.3231	-.3856	-.0624	18.0
WAT	-.2714	-.0830	-.1478	-.0648	18.7
Total (CM)	-	1.2637	0.9175	-.3462	100.0
			1976		
Constant	2.3738	2.3738	2.3738	-	-
MED	-.1972	-.1660	-.2674	-.1014	32.0
HED	-.2165	-.2016	-.3064	-.1048	33.1
TAX	-.1688	-.7736	-.8185	-.0449	14.2
HY	-.0051	-.0260	-.0311	.0050	-1.6
WAT	-.2954	-.0903	-.1609	-.0705	22.3
Total (CM)	-	1.1683	0.8517	-.3166	100.0

impact of the water variable, since women in households that had recently been linked to the water system could be reporting higher mortality from the period before they had access to the system.

Attention was focussed on both the level of child mortality and differences among the socioeconomic groups during the period covered by the two data sources. In the underlying conceptual framework, both household and community variables relate directly to child mortality as determinants in their own right as well as indirectly through variables that are endogenous at the household level, water being the case in point, either as demand variables (at the household level) or supply variables (at the community level). A recursive model consisting of three equations was estimated, and the results used to assess the contribution of increases in access to piped water in urban Brazil to declines in the level of child mortality by means of path analysis.

The main conclusions about these differentials are as follows:

1. That exogenous household variables (education of mothers and husbands) had the greatest total effect on differences in child mortality and that most of this effect was direct.
2. That access to piped water had a significant but secondary impact on differences in child mortality, accounting for about one-fifth of such differences.
3. That community level variables played a minor, but still significant role in child mortality differentials, with most of the effect being indirect via the household water access variable (WAT), which was expected since they entered the model via the water supply equation in the specification.

Examination of changes from 1970 to 1976 indicated that most of the decline in child mortality could be accounted for in terms of changes in the composition of the study population rather than changes in the parameters of the causal model, with increased education of both mothers and husbands playing a prime role, and increased access to piped water a secondary role. Some changes were observed in the parameter estimates, ones consistent with the expectation that increased water supply could contribute to attenuation of differentials in child mortality associated with education and income. The relative weight of supply variables in the decomposition of child mortality differentials increased from 1970 to 1976, though most of this increase continued to be indirect, via WAT, the household level water access variable. Interaction terms were introduced as a test of the strength of this attenuation effect, and indicated that in 1976 the effect of water was somewhat greater among lower income groups.

The ameliorating effect of increased access to piped water on income class differentials in early childhood mortality was thus significant but limited in comparison with other household level characteristics, particularly education. This does not imply that community level interventions were ineffective; though the level of education was specified as exogenous at the household level in this study, it was also the product of community level action and could have both community and individual level effects. Brazil chose to invest in education earlier than it chose to invest in water supply, and differences in the relative impact of the two variables may also reflect these differences in timing.

NOTES

[1] The analysis is limited to water supply even though the PLANASA program included both water and sanitation, because preliminary analysis indicated that in urban areas (where the program was being implemented) almost all households that had piped water also had a link to the city sanitation system, or to another "modern" sanitation system (e.g., septic tanks). Most of the cells required for testing their separate effects (water but no sanitation, or vice-versa) were empty. Care should be taken in interpreting the results of this study to avoid any implication that the contribution

of safe water to increased child survival is more or less than improved sanitation, since that trade-off was not tested.

[2] Restricting the study population to women in their twenties also involved a decision *not* to attempt to analyze cohort changes (i.e. comparison of results for women 20–29 in 1970 to those for women 26–34 in 1976). Again, the reason for this was the time reference problem, since the length of the reference period increases with the age of the mother.

[3] Brazil's national system of vital statistics was established in 1974, although data for the municipalities of state capitals are available for earlier dates. While vital statistics for São Paulo and a few other states are reasonably complete and consistent with estimates from the census, they still fall short of the reliability and completeness of coverage needed for analysis of national trends and differentials (Altmann and Ferreira, 1979, 55). A thorough assessment of the available data on Brazilian fertility and mortality during the period from 1950 to 1976 has been carried out by the National Academy of Sciences Committee on Population and Demography (1983). The report on Brazil provides indirect mortality estimates derived from questions on children ever born and children surviving as reported in the 1950, 1960, and 1970 censuses and in PNAD surveys taken after 1970 for Brazil as a whole and for each of the seven PNAD sampling regions (see Appendix Figure). The rural areas of frontier states (the North and Central-West macroregions, which accounted for about 5 percent of Brazil's population in 1970) were not included in the PNAD samples and are excluded from the NAS estimates. Results of the 1980 census have since been released, and it was also possible to calculate a national level estimate of the infant mortality rate from 1980 data.

[4] A puzzling feature of the estimates derived from surveys during the 1970s is the divergence of rates for 1974–75 from those for 1972, 1973, and 1976. The 1974–75 estimates were calculated from special tabulations of the special nutrition survey (ENDEF) conducted in those years. While the ENDEF results suggest a pattern of regional differentials that is similar to the ones in census and PNAD data, the level of infant mortality is significantly higher than in adjacent PNAD surveys. When ENDEF results are compared to the 1970 census, it appears that there was no decline at the national level, and that the Northeast/São Paulo differential increased even more than the PNAD indicates. The ENDEF results need to be viewed with caution, especially in comparison with estimates from the census and PNAD surveys. No assessment of the comparability of ENDEF to census and PNAD data is available. It is quite possible that ENDEF *is* closer to the Brazilian reality in terms of the level of mortality than the census or PNADs. However, this does not imply that mortality levels were constant after 1970; rather, it suggests that both the census and the PNADs may have understated the true level of mortality and that the decline occurred from a higher base.

[5] Though some of the "increase" in infant mortality rates could have resulted from improved reporting, the trends were associated with deteriorating conditions of housing and sanitation in the rapidly growing peripheral areas of larger cities (Leser, 1974; Monteiro, 1982), and with poor nutrition resulting from inflationary pressures on the limited purchasing capacity of low income families (Iunes et al., 1975). Data from the 1975 Estudo Nacional de Despesa Familiar (ENDEF) revealed that 37 percent of children under the age of 18 suffered from first-degree malnutrition, i.e. 76–90 percent of normal weight for age (Knight and Moran, 1981, 27–28).

[6] Published data from the 1976 PNAD survey include tabulations of child survival ratios by minimum wage level, but tests of estimates derived from these tabulations revealed problems of comparability with the categories employed by Carvalho and Wood (1978) for 1970, so the data were retabulated. Household income was calculated by adding up the monetary earnings of family members (after imputing earnings to members whose earnings were reported as unknown—see Fox, 1981). The 1976 survey included several types of income not explicitly covered in the 1970 census. To insure comparability, 1970 concepts have been applied in both instances. Boarders and household employees were excluded, while other related individuals are included, so that "household" income measures extended family income. Estimates are presented for both urban and rural women, though the latter are excluded from further analysis. Also excluded are women not related to the head of the household in which they resided, a step taken to eliminate bias arising from the inclusion of live-in domestic servants in a substantial proportion of middle and upper income Brazilian households.

[7] The life table values reported in the table were derived from child survival ratios in the census and PNAD data, which were adjusted using multipliers based on ratios of parity at ages 15–19 to parity at ages 20–24. Because of the limited number of births to higher income women at these ages, the parity ratios could be distorting the adjustment. The observed survival ratios and parity ratios are as follows:

Income deciles	P1/P2	Survival ratio
	1970	
1–3	.087	.86
4–6	.129	.86
7–9	.157	.88
10	.118	.93

Income deciles	P1/P2	Survival ratio
	1976	
1–3	.109	.88
4–6	.142	.89
7–9	.147	.93
10	.076	.97

The survival ratios suggest that the widening differential between the upper and lower groups does not derive from distortions in the adjustment factors, though the pattern of average parity differentials does suggest caution about the appropriateness of the assumptions of indirect measurement of life table values when populations are classified by income. One effect of grouping by income is that ${}_3q_0$ values for women in higher income deciles are lower rather than higher than ${}_2q_0$, which probably results from changes in the composition of groups as aging moves women into higher income deciles.

[8] Data on the relation between child mortality, nutrition, and access to health care in Brazil are limited to local area studies such as the ones reported by Puffer and Serrano (1973). The one exception is the 1974–75 ENDEF survey; however, the results of that survey are available only in the form of published summary tables for large regional aggregates, and the tabulations that have been published do not include any that link survival rates to nutrition levels or other aspects of household consumption that were included in the scope of the survey. Since a public use file of ENDEF data has not been released, special tabulations were limited to variables included in the 1970 census and 1976 PNAD survey, which are quite limited in comparison with ENDEF.

[9] "Supply" is used here in the more restrictive sense of microeconomic theory than in the usual reference to "water supply", which includes both supply *and* demand in the economic sense.

[10] Two additional problems remain in the use of OLS for these equations. One relates to measurement of the dependent variable, which as measured has statistical properties that make OLS an inappropriate estimating technique. A transformation of CM that yields the desired properties is discussed in the next section. The second concerns equation [2], whose dependent variable is WAT, a dummy variable taking values of zero and one depending on whether a household has water or not. With a dummy dependent variable, OLS can lead to predictions of values less than zero and greater than one, in addition to yielding inefficient estimates, so that logistic maximum-likelihood estimating techniques are preferred. However, the error term restriction described in the previous paragraph becomes much more complex when OLS is not used; further, OLS is least likely to manifest these properties when the sample mean of the variable to be estimated falls between 0.3 and 0.7, which is the case with these data (Amemiya, 1981, 1488). Therefore the decision was made to estimate equation [2] with OLS.

[11] The mortality standard that was employed for both data was the Coale-Demeny South model, level 19 ($e_0 = 65$). Average parity ratios (P_1/P_2 and P_2/P_3) were calculated separately for each date.

[12] Two assumptions required for application of the Trussell-Preston method are violated by the Brazilian data. These relate to geographic and interclass mobility within the study population and to changes in mortality levels in the years prior to the time that retrospective questions are asked. Since the index is based on the cumulative experience of women's births and surviving children, the mortality level observed at a particular point in time will not be representative of women in a particular reference group (for example, an income class) if there has recently been significant movement of women between classes, or if mortality rates have been declining rapidly. One way to reduce the time bias is to restrict the analysis to women in their twenties, whose experience is more recent. Table 2 suggested that the experience of younger women reflected mortality rates 2–3 years before the 1970 census and 1976 PNAD survey, whereas the time reference of rates for older women was 3–5 years. In dealing with the mobility problem, Carvalho and Wood (1978) have suggested the selection of standard schedules based on groups that are broad enough to minimize mobility effects. Following this suggestion, separate standards were used in calculating the index depending on whether women were members of families that were above or below the median level of family income.

[13] Zero-order correlations between error terms for the endogenous variables are as follows. Coefficients are shown above the diagonal and their significance (the likelihood that the correlation is random) below it.

	HY	WAT	CM
		1970	
HY	–	.00	.00
WAT	.99	–	.00
CM	.99	.99	–
		1976	
HY	–	.00	.07
WAT	.99	–	.00
CM	.47	.99	–

In all but one instance the zero order correlations between the error terms and exogenous variables are not significantly different from zero. The exception is MED and the error term for HY, owing to the high correlation between MED and HED.

MED is not included in the HY equation on the ground that this correlation is related more to the selectivity of the marriage process than to spouse contributions to earning capacity.

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Appendix Table.—Zero order correlation coefficients between variables in Table 5, 1970 and 1976[a]

	MED	HED	TAX	WS	HY	WAT	CM
MED	-	.589	.188	.189	.514	.389	-.205
HED	.673	-	.221	.227	.572	.421	-.215
TAX	.292	.341	-	.936	.339	.317	-.155
WS	.350	.419	.843	-	.312	.336	-.150
HY	.441	.488	.350	.378	-	.481	-.178
WAT	.368	.392	.331	.404	.312	-	-.196
CM	-.187	-.181	-.135	-.146	-.144	-.149	-
PWS[b]	-.058	-.060	-.186	-.317	-.009	.025	.010

[a]Coefficients for 1970 are shown above the diagonal; those for 1976 below it.

[b]PWS was measured in 1976 only.

Appendix Figure.