Bequest Behavior and the Effect of Heirs' Earnings: Testing the Altruistic Model of Bequests

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That parents transfer resources to children because of altruistic concern is a reasonable a priori assumption. However, economic theories of altruistic transfers have produced many counterintuitive conclusions, and, consequently, much debate. When applied to bequests, these theories predict that inheritances will compensate for earnings differences between siblings as well as between parents and children. This paper tests these implications. Using a new data set centered on federal estate tax returns, little support can be found for an altruistic theory of bequests. This finding has implications for macroeconomic policy, government transfer programs, and inequality. (*JEL* D19)

Hundreds of billions of dollars per year are bequeathed by people in the United States,[1] and about two thirds of this passes from parents to children.[2] Among several economic theories of such intergenerational transfers, the most prominent is that bequest behavior is motivated by altruism. As defined by Robert J. Barro (1974) or Gary S. Becker (1974), this means that parents bequeath because they gain utility from the utility or lifetime resources, respectively, of their children. It follows that bequests are compensatory. Parents will bequeath unequal amounts to their offspring, compensating children who have low earnings. In addition, bequests compensate intergenerational differences, with lower average earnings of children eliciting larger bequests from parents.

Although altruistic bequest models have raised much controversy in the theoretical literature, they have been subject to few empirical investigations. Using a new data set of parents' estate tax returns linked to the parents' and children's income tax returns this paper finds evidence which is generally inconsistent with the compensatory bequest implications of altruistic models. This lack of support for an altruistic bequest theory has implications for the relevance of Ricardian equivalence propositions, the degree to which private transfers may undo government redistributional efforts, the relationship between inheritance and inequality, and for the use of altruistic bequests to legitimize models with infinitely-lived agents.

The paper is organized as follows. The next section reviews the previous empirical work

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[1] The exact figure is unknown. Based on estate tax data discussed in Marvin Schwartz (1988) and Mary F. Bentz (1984), a rough calculation is that in 1982 about $200 billion was transferred at death. This is in line with Harry L. Gutman's (1979) also rough estimate of $131 billion transferred at death in 1977.

[2] Again, the exact amount inherited by children is unknown. The marital deduction is approximately one third of aggregate gross estates (see Bentz, 1984). From the smaller sample used in this paper, children's inheritances are twice the level of bequests to spouses. Combining these two figures leads to the rough estimate of two thirds. Robert B. Avery and Michael S. Rendall (1994) estimate that children inherited $39.4 billion in 1990, but note this is likely an underestimate.

on intergenerational transfers. The data are introduced in Section II. Section III discusses the altruistic models that are estimated. The results in Section IV are followed by concluding comments.

I. Previous Work

Data on heirs' incomes are not necessary for preliminary investigations of bequest practices. Therefore, several authors have studied the incidence of unequal division of estates to glean some indirect evidence of bequest behavior. From Connecticut probate records, Paul L. Menchik (1980) found that most parents bequeath equal amounts to their children. However, the predominance of equal division has been challenged by Nigel Tomes (1988) who used survey data from Cleveland to show that most parents divided their estates unequally. Menchik (1988) cast doubt on this result by replicating Tomes' sample using probate records and finding that most parents were equal dividers. Denis Kessler and Andre Masson (1988a) consider this empirical debate unresolved, noting the importance ascribed to unequal estate division in the U.S. legal literature, and, elsewhere (1988b), calling for additional empirical evidence. In addition to the potential survey response error in Tomes' data, the divergence in estate division results may be caused by the small sample sizes being considered.

The data requirements for bequest analysis beyond the division of estates are stringent, but nevertheless several researchers have made use of some data sets which contain information, albeit incomplete, on parents and children. B. Douglas Bernheim et al. (1985) found evidence in the Longitudinal Retirement History Survey that children visited and called their parents more frequently when their parents had larger amounts of bequeathable wealth. They interpret this finding as evidence of an exchange theory of transfers in which bequests are made to children in exchange for their earlier attention and care. Parents are able to elicit attention because they can threaten any child not providing attention with the credible promise to disinherit him in favor of his siblings. However, parents with only one child would be unable to credibly make this threat, and, consistent with this argument, the positive link between visits and bequeathable wealth was not found in one-child families. In contrast to this evidence supporting an exchange model, and using the only previously available bequest data which have included information on heirs' incomes, Tomes (1981, 1988) found evidence that bequests were compensatory both across families and within families, respectively. In addition, he found that visits by children were not related to the inheritances they later received.

Clearly bequests are not the only financial transfer between parents and children; in the aggregate, inter-vivos gifts are of the same order of magnitude.[3] Examination of inter-vivos gifts by Cox (1987) and Cox and Mark Rank (1992) have found that, conditional on the event of a positive transfer having occurred, inter-vivos transfers increased as the income of the recipient increased. Such a finding is consistent with a theory of transfers based on exchange between parents and children, but not with altruism.

Joseph G. Altonji et al. (1992) have tested consumption and income data for the implicit presence of compensatory inter-vivos transfers among extended family members. If transfers are altruistically motivated, within-family consumption differences should be independent of the within-family distribution of income, but the evidence is to the contrary. In addition, more general tests of the within-family income-consumption covariance restrictions implied by altruism have failed to support the model, although the rejection is slightly weaker for the subset of adult children who are expecting a bequest (Fumio Hayashi et al., 1991).

Thus, the available empirical research offers an inconsistent picture of the behavior behind bequests and inter-vivos gifts. Within the

[3] Aggregate estimates vary by data source. Data from the Survey of Consumer Finances (William G. Gale and John Karl Scholz, 1994) show bequests and inter-vivos gifts to be of approximately the same magnitude, but data from the President's Commission on Pension Policy (Donald Cox and Fredric Raines, 1985) indicate that the latter are twice as large as the former. Tax data imply that aggregate bequests are larger than either study suggests (see footnote 1).

bequest literature there are conflicting results, even regarding the seemingly straightforward issue of the frequency with which children inherit equal amounts. Note that if equal division is the rule, then exchange and altruistic explanations of bequest behavior are somewhat discredited because they both imply that bequests should vary according to the characteristics of children.

II. The Estate-Income Tax Match Data

A. *The Data*

The Estate-Income Tax Match (EITM) data set was constructed by the Statistics of Income Division of the Internal Revenue Service. The federal estate tax returns of 1982 decedents were merged with their own and their beneficiaries' (1980–1982) income tax returns. The initial subsample is a 1-percent random sample of decedents with gross estates over $0.3 million but under $1 million (453 decedents) and a 100-percent sample of gross estates exceeding $1 million (8056 decedents). David Joulfaian (1992) has described the data in more detail. These data have several advantages over those previously available for bequest analyses. Most significantly, they are the first bequest data which contain reliable income information for *both* parents and child heirs. In contrast, Tomes (1981) had no parental income data. The data set contains complete families (rather than only a single child per parent), indisputably matched together via social security numbers (a weak link in some intergenerational data sets). Finally, the data permit separation of labor earnings from returns to capital, do not suffer from survey response bias, are national in coverage, and contain substantially more observations than earlier studies.

Summary statistics for the 4188 decedents who bequeathed directly (as opposed to indirectly via trusts) to natural born or adopted children are presented in column 1 of Table 1.[4] The average gross estate among the decedents is $2.56 million and their net wealth is of similar magnitude. Not surprisingly, their average annual income ($169,400) was high.[5] Income tax returns from either 1980 or 1981 were found for 96.5 percent of the decedents. Slightly more than half of the decedents were survived by spouses, and they bequeathed to 2.26 children, on average. The average child inheritance is $238,200, and the average absolute value of the deviations of children's inheritances from within-family averages is $14,600. Note that this reflects many children who receive bequests equal to those of their siblings. Forty-five percent of these decedents had children who are of different sexes. The decedents' average age was 75 years and only 39.5 percent were female. One quarter used trusts to transfer part of their estate. Twenty-five percent made inter-vivos gifts during the three years prior to death; these amounts are included in the gross estate.

It was more difficult to match income tax returns to the child beneficiaries. At least one return from 1980–1982 could be found for 73.2 percent of the children named as heirs on the estate tax returns.[6] Descriptive statistics for these matched children are presented in column 2, as are statistics for their parents, likewise conditional on being matched. The average age of the children is 42.7 years. Forty-nine percent are female and 66 percent are married. On average, the children report 1.092 dependents, likely an underestimate of

[4] Evidence (available from the author upon request in Appendix A) implies that these decedents were neither disproportionally less wealthy nor younger. A probit model of the probability of bequeathing directly to a child shows positive, but insignificant, relationships to both the size of the gross estate and the age of the decedent.

[5] Decedent income is defined to be wages and salaries, earnings from self-employment and partnerships, and returns from capital. The positive income concept (negative income components are set equal to zero) is used to reduce the noise caused by tax-shelter behavior (see Susan Nelson, 1985). When returns from both 1980 and 1981 were found, income is the two-year average.

[6] If children's income tax returns were not found because they had low income, and if they received larger inheritances, then excluding these unmatched children would stack the data against the altruistic model. Together these hypotheses imply that the larger a child's inheritance, the less likely would the child's tax return be found. In contrast, Appendix A (available from the author) indicates that the probability of matching a child's tax return increases with his inheritance.

TABLE 1—DESCRIPTIVE STATISTICS

	Decendents with one or more children (Sample 1)	Matched children (Sample 2)	Decedents/children matched; restricted (Sample 3)	Completely matched multichild families (Sample 4)
Parents (decedents):				
Gross estate	25.635	26.350	25.498	27.661
	(77.179)	(88.588)	(97.803)	(132.110)
Net wealth	23.946	24.676	24.108	26.391
	(73.630)	(84.386)	(92.758)	(125.210)
Parent income	1.694[a]	1.723	1.671	1.710
	(2.789)	(2.926)	(2.584)	(3.118)
Parent age	75.347	75.627	76.982	77.724
	(12.104)	(11.484)	(9.782)	(8.831)
Parent sex	0.395	0.390	0.409	0.425
	(0.489)	(0.487)	(0.491)	(0.494)
Surviving spouse	0.519	0.524	0.496	0.489
	(0.499)	(0.499)	(0.500)	(0.500)
Number of children	2.259	2.281	2.171	2.476
	(1.276)	(1.273)	(1.142)	(0.756)
Children are of different sexes	0.451	0.453	0.440	0.599
	(0.497)	(0.497)	(0.496)	(0.490)
Trust	0.256	0.251	0.264	0.239
	(0.436)	(0.433)	(0.441)	(0.426)
Lifetime gifts	0.253	0.252	0.257	0.264
	(0.435)	(0.434)	(0.437)	(0.441)
Families (number of decedents)	4188	3010	2020	948
Children:				
Inheritance	2.382	2.468	2.690	2.561
	(2.989)	(2.983)	(2.898)	(2.397)
Earnings	—	0.461	0.530	0.552
		(0.915)	(0.771)	(0.753)
Age	—	42.685[b]	44.752	45.564
		(12.432)	(10.016)	(9.402)
Sex	0.507	0.494	0.488	0.488
	(0.499)	(0.500)	(0.499)	(0.499)
Spouse	—	0.661	0.743	0.768
		(0.473)	(0.436)	(0.421)
Grandchild (children's children)	—	1.092	1.223	1.297
		(1.289)	(1.300)	(1.320)
Average inheritance of children[c]	0.146	0.137	0.142	0.155
	(0.592)	(0.557)	(0.593)	(0.646)
Average earnings of children[c]	—	—	—	0.272
				(0.415)
Number of children:	9464	6928	4153	2348
Used in tables:	2	—	3, 5	4, 5

Notes: Dollar amounts are in $100,000's (1982 dollars). Entries are omitted when their interpretation is ambiguous: that is, children's information when the sample contains unmatched children; and intersibling differences when the sample contains unmatched siblings.

[a] Income averaged over the 4042 matched decedents.

[b] Averaged over the 6212 children who had nonmissing age data.

[c] Absolute value of deviation from within-family average.

the decedents' grandchildren because many grandchildren may have already formed independent households.

The average inheritance of matched children is $246,800, over five times their average labor earnings of $46,100. Note that, on the basis of their labor earnings alone, the average child beneficiary is well above the lower bound ($39,704) of the top quintile of the family income distribution. Because labor earnings are intended to proxy children's human capital, they are defined to be the sum of wages and salaries and earnings from self-employment. Partnership income is added only if it dominates each of these two components. Otherwise, it is likely that the partnership is being used as a tax shelter, a common practice among high income individuals (Nelson, 1985). In addition, a lower bound of zero is assumed for each income component. If the children are married, the combined earnings of spouses are used, and bequests to such children are combined with bequests given to their spouses. Finally, each child's earnings are averaged over as many years between 1980–1982 as could be matched. This reduces the influence of the annual transitory component of earnings.[7]

In addition to the requirements that children be matched to an income tax return and that their parents be similarly matched, two sample restrictions are used when estimating the bequest models to be formulated below. First, children are excluded if they are less than twenty-five or greater than sixty-five years old (or if their age data are missing). This retains a focus on adult offspring and avoids complications surrounding transitions into and out of the labor force. Second, children from farm families are omitted due to the difficulties in using tax data to deduce the economic income from farms. Each of these restrictions removes nearly 1300 children from the sample.[8] The resulting sample contains 4153 children; their summary statistics are listed in column 3. As expected from the restrictions, these children are somewhat older, more likely to be married, and have higher earnings than those in column 2.

Some of the models considered below include the decedent's decision to divide her estate among several children. The data used to estimate these models contain additional exclusions. First, single-child families are excluded because there is no decision concerning the division of the estate among children. Second, families are dropped from the sample unless income tax returns could be found for each child named in the estate tax return. This is done to avoid error in the measurement of sibling averages and within-family differences. Third, families are excluded if they contain any children less than twenty-five or greater than sixty-five years old. These exclusions reduce the sample of children to 2348 observations (column 4), primarily because of the requirement that families be completely matched.[9] In addition to the children's individual data, column 4 shows the average absolute value of the differences between children's earnings and the within-family means to be $27,200.

Several limitations of these data should be noted. First, observations of children's inheritances are those amounts received directly and do not include assets transferred in trust. Unfortunately, the life tenants (recipients of income flows from such assets) and remaindermen (eventual recipients of such assets) of bequests via trust are unidentified in the data. Hence, the amounts inherited directly by the children of the 25 percent of decedents who created trusts may simply be a lower bound to their total inheritance. Though the results presented below refer to bequests made directly to children, as part of a sensitivity analysis the findings will be reexamined with the decedents forming trusts removed from the sample. Also note that although some trusts are certainly established for children with health problems which cause low earnings, Menchik's (1980)

[7] Eighty-one percent of all matched children were matched in each of the three years. Thirteen percent were matched in two years and 6 percent in only one year.

[8] Nearly 200 children are dropped because their parents were not matched to an income tax return. Eleven children are dropped because either their parent reported zero income in the two years prior to death or made a zero direct bequest to them.

[9] Sample 4 includes one family (two children) in which the decedent reported zero income. This observation is included in the within-family analysis, but is dropped from the between-family analysis.

evidence implies that most provide equal remainder interests.

Second, if compensatory inter-vivos transfers are frequently made and of substantial magnitude, it is possible that altruistic parents do not need to make compensatory bequests. Because the EITM data do not inform us of the inter-vivos transfers made from a parent to her children, except for those made during the last three years of her life, this possibility cannot be definitively ruled out. However, there are several indications that extensive compensatory inter-vivos gifts are unlikely. Recall, that previously discussed inter-vivos gifts research generally produced results inconsistent with altruism. Also, the EITM data suggest that inter-vivos gifts were not frequently made by the parents; only 25 percent made such gifts in the three years prior to death. A more conservative interpretation of this result is that it reflects a hesitancy to make inter-vivos transfers on a regular basis, not an unwillingness to make gifts when children experience times of reasonable need. Note that because the average EITM child was in his forties when his parent died, the major occasions of such need (for instance, help with a down payment on a first house and assistance at the birth of children) would have occurred well in advance of the parent's death. Consequently, gifts in response to such need would not be observed in the data. Even so, EITM children received an average inheritance of nearly \$250,000; achieving a transfer of similar magnitude with inter-vivos gifts would have required a sustained level of large gifts in addition to the transfers made at the times of need. Parents are likely aware of the deleterious effects such large inter-vivos gifts may have on their children's behavior (see Neil Bruce and Michael Waldman, 1990).

The third data issue is whether the number of children in a family can be accurately determined from the EITM data because only children receiving direct inheritances are required to be recorded on estate tax returns. Living children would not be observed in the data if their entire inheritance was received as a trust transfer or if they had been disinherited. Similarly, and unfortunately, stepchildren were coded into a residual relationship category, and hence cannot be identified. If these problems lead to the existence of many children who are not observed in the data, then the observed family size distribution should be skewed toward smaller families. Appendix B (available from the author upon request) presents a comparison of the family-size distributions in the EITM data with those from other data sets, and shows that EITM family sizes are in fact smaller. To the extent that this reflects the omission of disinherited children, and if this disinheritance is motivated by altruistic parents choosing bequests at corner solutions, then the analysis will be biased against compensatory bequests. However, other data sets have found that disinheritance occurs infrequently (Marvin B. Sussman et al., 1970; Menchik, 1980) and usually for reasons consistent with an exchange motivation (Sussman et al.). Moreover, the EITM data are drawn from the extreme upper tail of the wealth distribution, and there is a inverse relationship between socioeconomic status and fertility (Judith Blake, 1989). Hence, although there is undoubtedly some unobserved disinheritance, the many small families in the data do not necessarily indicate extensive disinheritance. All of these arguments are more fully developed in Appendix B. Despite these indications that disinheritance is not widespread, its incidence in the EITM data cannot be directly ascertained, and this paper's results should be interpreted with this in mind.

Finally, note that these data describe only the wealthiest of decedents, and therefore are not representative of the U.S. adult population. Studying approximately the same sample from which the EITM data were drawn, Schwartz (1988) estimated that 1982 decedents who filed estate tax returns represented the richest 2.8 percent of adults who in turn owned about 30 percent of U.S. personal wealth.[10] Although this percentage is not precisely accurate,[11] it is

[10] The EITM decedents were subject to a slightly lower gross estate filing threshold (Schwartz used \$325,000) and were necessarily matched to their income tax returns. As indicated above, this match did not impose much selection.

[11] The estimate is biased because it was based on mortality rates which were not adjusted for education, occupation, and other individual characteristics which affect mortality. For instance, Martin H. David and Menchik (1988) found that estate multiplier estimates of wealth are too low if not corrected for occupation.

TABLE 2—INCIDENCE OF EQUAL DIVISION AND INHERITANCE INEQUALITY AMONG CHILDREN

	Number of children					
	Two or more	Two children	Three children	Four children	Five children	Six or more
Estate division:						
Exactly equal	0.686	0.697	0.700	0.651	0.672	0.555
Within ± 2 percent	0.766	0.776	0.782	0.718	0.793	0.616
Inheritance inequality:						
Coefficient of variation (squared)	1.3308	0.9940	1.4123	1.2649	1.1147	1.0834
Within-family component	0.1007	0.0927	0.0569	0.1541	0.1763	0.1347
Families:	2913	1531	797	370	116	99

Note: The sample is all decedents in Sample 1 of Table 1 who bequeathed to more than one child.

a qualitative indication that the EITM data are representative of a substantial share of individuals' wealth, despite their representation of only a small percentage of individuals.

B. *Equal and Unequal Division Among Children*

Table 2 reports the incidence of equal estate division in the EITM multichild families (Sample 1 with single-child families excluded). Over two thirds (68.6 percent) of the decedents divided their estates exactly equally among their children. Over three quarters (76.6 percent) divided their estates so that each child received within ±2 percent of the average inheritance among children in the family.[12] Eighty-eight percent divided their estates "approximately equally,"[13] compared to 50.4 percent in Tomes (1988).

The incidence of equal division of estates is extremely high. The frequency of exactly equal estate division was reported to be 21.1 percent by Tomes (1988), 62.5 percent by Menchik (1980), and 84.3 percent in Menchik (1988). The table shows no systematic relationship between division choices and family size, except that very large families experience more unequal division. Finally, the squared coefficient of variation of inheritances received by children indicates substantial inheritance inequality. However, the within-family component of this inequality measure shows that very little inequality results from unequal inheritances to children within the same family.

III. The Altruistic Bequest Model

In this section the standard altruistic model is generalized to be consistent with the extensive amount of equal division which appears in the data. To begin, consider the standard model in which the utility of the jth parent, $U(c_{0j}, y_{1j}, y_{2j}, \ldots, y_{N_j j})$, is defined over her own lifetime consumption (c_{0j}) and her children's lifetime resources (y_{ij}). The subscript "0" indicates a parent variable, and $i = 1, \ldots, N_j$ indexes her natural born and adopted

[12] Unequal bequests that occur via trusts cannot be observed in the data. Indeed, decedents creating trusts are less likely to bequeath equal amounts (63.2 percent) compared to decedents not creating trusts (70.2 percent). Of course, unequal bequests may be made via trusts, despite equal direct bequests, or unequal direct bequests may be rendered inconsequential by very large and equal bequests via trusts. A referee points out that while children generally receive equal shares of remaining interests (excepting larger amounts for children with atypical disadvantages), creating generation-skipping trusts of equal amounts in which children receive a life interest are essentially unequal bequests because the children have different lengths of life. Making the very conservative assumption that every trust creator bequeathed unequally, then the true incidence of equal division among all the EITM decedents would be 54.3 percent. Note that Menchik's (1980) equal division results do include amounts transferred via trust.

[13] Tomes (1988) defined "approximately equal" estate division to be when the difference between the maximum and minimum sibling inheritances is no more than one quarter of the mean inheritance per child.

children, the number (N_j) of which is assumed exogenous. As is regularly assumed, the parent has symmetric concern over her children; the y_{ij} enter the utility function symmetrically. The lifetime resources of children are defined to be

$$(1) \qquad y_{ij} \equiv h_{ij} + b_{ij}$$

where b_{ij} is the bequest from the parent and h_{ij} is the exogenous lifetime earnings (human capital) of the child. The parent then chooses c_{0j} and the b_{ij} to maximize utility subject to a budget constraint determined by her lifetime resources (y_{0j}) and the intergenerational discount factor ρ. It is assumed that parents take ρ as fixed. It is convenient to reformulate the choice problem in terms of y_{ij} instead of b_{ij} (Becker, 1981), yielding the budget constraint:

$$(2) \qquad c_{0j} + \rho \sum_{i=1}^{N_j} y_{ij} = y_{0j} + \rho \sum_{i=1}^{N_j} h_{ij}$$

where the right-hand side is "family income," a family-level variable henceforth denoted Y_j.

The first-order conditions produce a solution, $f(Y_j, \rho)$, for children's lifetime resources which is a function of family income and the discount factor, and which with (1) yields a prediction concerning bequests in standard altruistic models:

$$(3) \qquad b_{ij}^* = -h_{ij} + f(Y_j, \rho).$$

The assumption of symmetric concern implies that the scalar $f(Y_j, \rho)$ does not vary across children in the same family. Hence, when control-ling for this family-specific scalar, altruistic bequests are a negative linear function of earnings.

A stochastic specification which is similar to (3) can be developed if utility is modified to be a function of $c_{0j} - \gamma_{0j}$ and $y_{ij} - \gamma_{ij}$, where γ_{0j} and γ_{ij} are minimum demand levels for consumption and children's lifetime resources, respectively. Assume the minimum demand levels can be written

$$(4a) \qquad \gamma_{0j} = \mathbf{D}_{0j}'\boldsymbol{\delta}_p + \eta_j$$

$$(4b) \qquad \gamma_{ij} = \mathbf{D}_{ij}'\boldsymbol{\delta}_c + \varepsilon_{ij}$$

where $\mathbf{D}_{0j}$ is a vector of parent-specific demographic characteristics, $\mathbf{D}_{ij}$ is a vector of child-specific demographic characteristics, $\boldsymbol{\delta}_p$ and $\boldsymbol{\delta}_c$ are the respective parameter vectors, η_j is a parent-specific heterogeneity term, and ε_{ij} is a child-specific heterogeneity term (for example, it would include unobserved child needs). Each error term is independent and identically distributed. Note that although (4b) allows bequests to be affected by observable and unobservable characteristics of children, assuming these characteristics influence bequests independent of the identities of the children possessing them implies that utility remains symmetric in the $(y_{ij} - \gamma_{ij})$. With (4a) and (4b) optimal bequests are

$$(5) \qquad b_{ij}^* = -h_{ij} + \mathbf{D}_{ij}'\boldsymbol{\delta}_c + f(\hat{Y}_j, \rho) + \varepsilon_{ij}$$

where $\hat{Y}_j$ is family income redefined to include γ_{0j} and within-family sums of the γ_{ij}. Family income thus redefined remains a family-level variable; all of its stochastic components are family-specific unobservables. Symmetric concern again implies that the scalar $f(\hat{Y}_j, \rho)$ is common to all children within each family. Hence, $f(\hat{Y}_j, \rho)$ can be modeled as a family fixed effect in the estimation of (5); controlling for the fixed effects, the estimated coefficient on heirs' earnings enables a test of the altruistic model's prediction that this coefficient is -1. Moreover, note that imperfect observation of parental resources does not bias the test because family fixed effects control for y_{0j}.

However, there is a complication in applying this test to the data at hand. To see this, it is helpful to rewrite (5) in terms of deviations from within-family means:

$$(6) \qquad b_{ij}^* - \bar{b}_{\cdot j}^* = \beta_h(h_{ij} - \bar{h}_{\cdot j}) + (\mathbf{D}_{ij} - \bar{\mathbf{D}}_{\cdot j})'\boldsymbol{\delta}_c + (\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})$$

where the overbar denotes within-family averages and β_h is the degree to which parents use bequests to compensate intersibling differences in earnings (in the altruistic model $\beta_h = -1$). It is clear from (6) that equal division ($b_{ij}^* = \bar{b}_{\cdot j}^*$) should occur only in the unlikely event that siblings have equivalent lifetime earnings and other characteristics. Because this contradicts the prevalence of equal division documented above, a generalization of the standard model is required. I generalize the model by assuming that the parent suffers a psychic cost κ_j, resulting from intersibling jealousy and family conflict, if she chooses to divide her estate unequally among her children. This seems entirely plausible, although other explanations of equal division exist.[14] The decision process in this model can be thought of as having three steps. First, the parent determines the optimal unequal bequests as in (5). Next, she calculates the optimal equal bequests (b_j^{**}) which maximize utility subject to (2) and the additional constraint that $b_{ij} = b_{kj}$ for all children i and k in her family. Finally, she selects unequal division only if it gives her higher utility after the psychic cost of unequal division is deducted, that is if $U_j^* - \kappa_j > U_j^{**}$.

This expression is an estate division rule. Using a second-order Taylor series expansion of U_j^{**} around the optimal unequal bequests b_{ij}^*, and assuming that parental utility is separable in the $y_{ij} - \gamma_{ij}$, leads to the following approximation of $U_j^* - U_j^{**} > \kappa_j$:[15]

$$(7)\quad b_{ij} = b_{ij}^* \quad \text{if} \quad N_j^{-1} \sum_{i=1}^{N_j} \{\beta_h(h_{ij} - \bar{h}_{\cdot j}) + (\mathbf{D}_{ij} - \bar{\mathbf{D}}_{\cdot j})'\boldsymbol{\delta}_c + (\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})\}^2 > \kappa_j$$

$$b_{ij} = b_j^{**} \quad \text{otherwise (for all } i \text{ in family } j).$$

The intuition of the division rule is straightforward. The likelihood of an unequal division is higher if siblings differ greatly in their earnings and other characteristics, including the unobservable ε_{ij}'s, and if parental psychic cost is low.

The generalized model recognizes that parents select themselves into the set of unequal dividers with a decision rule based on the same observable and unobservable child characteristics which determine the within-family bequest differences. Hence, it suggests a potential selectivity bias in estimates of within-family bequest models like (5) or (6) based on the subsample of families with unequal bequests. To illustrate this bias, consider using data from families with unequal divisions to run a regression of $(b_{ij} - \bar{b}_{\cdot j})$ on $(h_{ij} - \bar{h}_{\cdot j})$; also assume that altruism is the true model ($\beta_h = -1$). Then, equation (7) implies that in the sample of families for which unequal divisions are observed, it must be that siblings with greater [lower] than average earnings ($h_{ij} >$ [$<$] $\bar{h}_{\cdot j}$) tend to have lower [greater] than average unobserved components of minimal demand ($\varepsilon_{ij} <$ [$>$] $\bar{\varepsilon}_{\cdot j}$). If not (for example, $h_{ij} > \bar{h}_{\cdot j}$ and $\varepsilon_{ij} > \bar{\varepsilon}_{\cdot j}$) then the terms $-(h_{ij} - \bar{h}_{\cdot j})$ and $(\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})$ would offset each other in (7) and unequal division would be less likely. The implication is that $(h_{ij} - \bar{h}_{\cdot j})$ and $(\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})$ are negatively correlated within the subsample of unequal dividers. Consequently, regressing $(b_{ij} - \bar{b}_{\cdot j})$ on $(h_{ij} - \bar{h}_{\cdot j})$ without controlling for this negative correlation would produce estimates of β_h that are biased away from zero—that is, less than -1. In a multiple regression model, estimates of $\boldsymbol{\delta}_c$ are potentially biased as well; of course, in this case the covariance of earnings with the

[14] Menchik (1988) considers the psychic cost explanation more reasonable than a purely financial cost (of making a will) rationale, or explanations based on uncertainty over future generations' needs or survivability. Other possibilities are that parents are concerned with their children's opportunity sets (rather than their outcomes), or choose equal division to mitigate their children's strategic behavior. In the latter, equal division guarantees that kindnesses shown to parents by their children are motivated (at least at the margin) by the children's concern for the parent and not financial gain.

[15] The derivation of (7) is available upon request from the author in Appendix E. The separability assumption implies that the second-order cross-partial derivatives of utility are zero, and this leads to a Taylor expansion without interactions between the characteristics of different siblings. The derivation also requires that $b_j^{**} \approx \bar{b}_j^*$, that is, the average bequests to children are the nearly the same regardless of the estate division choice. However, this is simply a requirement that the second-order Taylor expansion be a good local approximation to U_j^{**} around the b_{ij}^*, as can be seen by noting that $b_j^{**} = \bar{b}_j^*$ if parental utility is quadratic. On average, the EITM data satisfy this requirement. In Sample 4, the average bequest to children receiving equally divided estates is $256,874; on average, children receiving unequally divided estates inherited $254,841.

$\mathbf{D}_{ij}$ prevents an a priori determination of the direction of any of the biases.

Unbiased estimates of β_h and $\boldsymbol{\delta}_c$ can be obtained by estimating (6) along with (7) as a generalized tobit model.[16] To implement this model, I will assume a linear specification of parental psychic costs:

$$(8) \qquad \kappa_j = \mathbf{Z}_j'\boldsymbol{\alpha} + \omega_j$$

where $\mathbf{Z}_j$ is a vector of parental characteristics ($\boldsymbol{\alpha}$ is the corresponding parameter vector), and ω_j is an i.i.d. $N(0, \sigma_\omega^2)$ unobservable component. Note that both the unobserved component and the total psychic costs may be negative. Therefore, the sensitivity of the results will be examined by considering a log-normal heterogeneity component and psychic costs which are the square of the right-hand side of (8).

Computing the contribution to the generalized tobit likelihood function from a family of N_j children requires either numerical integration of $(N_j - 1)$-variate densities over regions of the error space defined by the nonlinear relationship among error terms which results from (7) and (8), or using simulated maximum likelihood (SML). I estimate the model using SML with a kernel-smoothed frequency simulator (see, for example, John Geweke et al., 1994). The probabilities of equal and unequal division based on (7) and (8) are estimated by taking 50 draws for each of the $(\varepsilon_{ij} - \bar{\varepsilon}_{.j})$ and ω_j from their respective distributions; the $(\varepsilon_{ij} - \bar{\varepsilon}_{.j})$ are assumed to be $N(0, \sigma_\Delta^2)$. A disadvantage of simulated maximum likelihood estimation is that the simulated likelihood is a biased estimate of the true likelihood for a finite number of draws. Therefore, I have verified the performance accuracy of this procedure in a series of Monte Carlo experiments; these results are available upon request.

Finally, note that despite the potential selectivity bias, a test of the altruistic model based on estimating (5) with fixed effects is nevertheless of interest. First, if parental psychic costs dominate intersibling differences as determinants of unequal division, the bias will be small because the division decision is only weakly correlated with differences in children's characteristics.[17] Second, the bias may turn out to be weaker in families with more than two children. This is because when a parent chooses unequal division in a large family, it may be that just a few of her children have characteristics sufficiently far from the within-family averages to cause the left-hand side of (7) to exceed the psychic costs. The $(h_{ij} - \bar{h}_{.j})$ and $(\mathbf{D}_{ij} - \bar{\mathbf{D}}_{.j})$ are not correlated with the $(\varepsilon_{ij} - \bar{\varepsilon}_{.j})$ for the children who were *not* influential in their parent's division decision. Conversely, there is only one intersibling difference in two-child families and it necessarily influences the division decision (hence, the bias is likely to be largest in two-child families). Third, equation (5) is estimable with all matched children (Sample 3) without measurement error bias caused by the unobserved data of unmatched siblings. In contrast, the left-hand side of (7) cannot be accurately measured for families with one or more unmatched children. Hence, the generalized tobit model of (6) and (7) is estimated using the data from Sample 4.

Regardless of these differences, there are several advantages to using either the fixed-effects estimation of (5) or the generalized tobit estimation of (6) and (7) to test for an inverse relationship between bequests and earnings. Essentially, equations (5) and (6) result from a bequest model in which parents are the only decision makers (rather than being engaged in strategic exchange), children's lifetime resources less the minimum demand levels are the appropriate arguments in parents' utility functions (rather than the bequest amounts themselves), and parental utility is a

[16] A two-stage procedure (following James Heckman, 1976) cannot be used. Although an estimate of the magnitude of $E[(\varepsilon_{ij} - \bar{\varepsilon}_{.j}) \mid U_j^* - U_j^{**} > \kappa_j]$ can be formed from a discrete choice model based on (7), its sign is indeterminate because $(\varepsilon_{ij} - \bar{\varepsilon}_{.j})$ enters (7) in quadratic form only.

[17] The bias is also small in the reverse case where intersibling differences dominate psychic costs in the estate division decision. In this event there is little selectivity bias because there is essentially no selection. That is, among parents whose children's characteristics differ, almost all bequeath unequal amounts. As described in Section II, the EITM data do not reflect this pattern.

TABLE 3—INHERITANCE MODELS WITH FAMILY FIXED EFFECTS (UNEQUALLY DIVIDED ESTATES)

	Dependent variable: Inheritance		
Variables	(1)	(2)	(3)
Earnings	−0.12682[a] (0.07550)	−0.15135[a] (0.07896)	−0.17225[a] (0.09774)
Age	—	−0.02456[a] (0.01429)	−0.03081[a] (0.01838)
Sex	—	−0.01559 (0.1211)	−0.01513 (0.1584)
Spouse	—	0.29003[a] (0.1589)	0.35685[a] (0.2089)
Grandchild	—	−0.02347 (0.05457)	−0.02380 (0.06946)
Children:	1089	1089	821
Families:	527	527	396
Adjusted R^2:	0.625	0.626	0.497
F statistic:	2.821	1.815	1.653
(Probability value)	(0.089)	(0.107)	(0.145)

Notes: Observations are the children in Sample 3 who received unequal bequests. Family fixed effects are included in each regression. Earnings and inheritances are in $100,000's. Standard errors are in parentheses. The F statistic tests the hypothesis that all the independent variables, except the fixed effects, have zero coefficients. Likelihood ratio tests in column (1) indicate significance at 2 percent and in columns (2) and (3) at 1 percent. Columns (1) and (2) employ a strict definition of unequal division. Column (3) considers unequal division exceeding ± 2 percent of the average bequest to children.

[a] Significant at 10 percent.

symmetric function of these arguments (a standard assumption). Consequently, other than symmetry, tests based on (5) and (6) do not depend on functional form assumptions concerning preferences, that is, $f(\cdot, \cdot)$, which may be parent specific.

IV. Results

A. *Fixed-Effects Estimates*

Table 3 presents fixed-effects estimates of equation (5) using the children in Sample 3 whose parents chose unequal division. Assuming that earnings are the only child characteristics which affect bequests, the estimate in column (1) indicates that when a child's earnings are $1 below the within-family average child earnings, he receives almost $0.13 more than the average inheritance of that family's children. This compensation is small and, although significantly different from zero (at the 10-percent level), the 99-percent confidence interval (−0.32, +0.07) implies rejection of the altruistic null hypothesis that the coefficient is −1 at the highest levels of significance. Also note that (5) is an intergenerational model based on lifetime earnings, but it is being estimated with (a three-year average of) annual earnings. Adjusting the point estimate to account for the conversion of annual earnings into a lifetime stock variable implies substantially less compensation regardless of assumptions concerning interest rates and time horizons. For instance, an annual interest rate of 5 percent over an infinite horizon implies compensation of 0.63 cents on the dollar.

The estimated compensation is similar when the available demographic controls are added in column (2). These results also show that younger and married children receive more than their siblings. In column (3) the definition of unequal division is relaxed by defining it to have occurred if any child inherits more or less than 2 percent of the average bequest to children. The estimated effect of own earnings is slightly stronger.

Additional sensitivity analyses produce qualitatively similar results, the most interesting of which are now briefly described.[18] There is evidence that those who leave trusts simultaneously use direct bequests to children to compensate for earnings differences (coefficient = −0.734, standard error = 0.165). Of course, if the present value of bequests via trusts could be assigned to the appropriate children, there may be evidence of a stronger altruistic effect, but the effect is not likely to be large.[19] There is weak evidence that those not leaving trusts reinforce earnings differences. In addition, excluding decedents who made gifts during the years before death does not substantially change the estimates, and thus indicates that the behavior of decedents who may have already achieved their desired compensation during life is not driving the results.[20] Finally, intersibling compensation is larger when the decedent is either survived by a spouse (coefficient = −0.317, standard error = 0.109), has exactly two children (coefficient = −0.343, standard error = 0.194), or is among the 30 who had a gross estate over $5 million and bequeathed unequal amounts (coefficient = −1.308, standard error = 0.353). The first of these results suggests that for two parents to come to an agreement on an unequal treatment of their children, the children must have different needs.[21]

B. *Generalized Tobit Estimates*

Generalized tobit estimates of (6) and (7) are presented in Table 4. The variables used in this model require data on all of the children in each family; thus, Sample 4 is used rather than Sample 3. Ordinary least-squares (OLS) estimates of (6) using Sample 4 are presented for purposes of comparison. One observation per family is dropped to adjust for the loss of degrees of freedom by taking within-family differences. The estimates of the psychic cost parameters are relative to σ^2_ω, which has been normalized to 1. Columns (1) and (2) present results using the exact definition of unequal division; columns (3) and (4) use the "±2 percent" definition.

The OLS estimate of intersibling compensation in column (1) is −0.163 and is significant at 12 percent. That this is only slightly higher than the comparable fixed-effects estimate in Table 3 suggests little bias in moving from Sample 3 to Sample 4. The estimate falls to −0.126 in the generalized tobit model in column (2). Also, there is a decrease in the magnitudes of the point estimates on all the other child characteristics, except for the number of grandchildren. The estimates of the psychic cost parameters indicate a significantly positive and large average cost to unequal division, but none of the other parameters are significant. A chi-squared test of the restriction

[18] Appendix C (available from the author upon request) contains the full set of these results as well as subsequent results described in the paper but not included in Tables 3, 4, and 5.

[19] Only 25 percent of the parents in this study created trusts at their death, and only 10 percent of the distributable estates in the sample were transferred through trusts. Clearly, not all bequests in trusts go to children. Gerald Jantscher (1967) reported Treasury data on gross estates exceeding $0.3 million (equivalent to about $1.0 million in 1982 dollars) from the 1950's which show that children were involved as life tenants or remaindermen of trusts whose value equaled about 44 percent of all wealth bequeathed in trusts (unfortunately, more recent evidence is unavailable). Moreover, among those trusts that do pass to children it is unlikely that all of them affect a compensation for children's earnings differences.

[20] Also, in a probit model such giving is positively associated with choosing unequal division. To the extent that giving in the years preceding death indicates a greater propensity to make inter-vivos transfers in general, this result is inconsistent with the possibility that all necessary compensation is accomplished with inter-vivos giving.

[21] Estimating the model with subsamples defined by family sizes of three, four, and five or more children, other sizes of the parent's gross estate, the absence of a surviving spouse, and for the children who did not earn income from partnerships produces weaker evidence of within-family compensation than reported in Table 3. Estimating the model without the restrictions on the age of children yields an earnings coefficient slightly smaller than that in column (2). Although the altruistic model with symmetric parental utility unambiguously implies that inheritance is a linear (negative) function of own earnings, estimates of cubic, double-log, and within-family share functional forms do not produce stronger evidence of altruism.

TABLE 4—GENERALIZED TOBIT MODELS OF WITHIN-FAMILY INHERITANCE

Variables	Exactly equal division		Division within 2 percent	
	OLS (1)	Generalized tobit (2)	OLS (3)	Generalized tobit (4)
Children's characteristics:[b]				
Earnings	−0.16300	−0.12594[a]	−0.19160	−0.14326
	(0.10356)	(0.06797)	(0.13152)	(0.12291)
Age	−0.03535[a]	−0.01744[a]	−0.04428[a]	−0.01698
	(0.01864)	(0.00942)	(0.02407)	(0.01123)
Sex	−0.05238	−0.03936	−0.06350	−0.02916
	(0.14462)	(0.07401)	(0.19105)	(0.08773)
Spouse	0.27877	0.16492[a]	0.36753	0.16772
	(0.18355)	(0.09572)	(0.24590)	(0.11209)
Grandchild	−0.00317	0.00768	0.00002	0.00436
	(0.06442)	(0.03224)	(0.08417)	(0.04015)
σ_Δ	1.156	0.816	1.345	0.862
Parent's psychic cost:				
Parent income	—	−0.01208	—	−0.02096*
		(0.01131)		(0.00903)
Surviving spouse	—	−0.04897	—	0.23538[a]
		(0.10491)		(0.12821)
Number of children	—	−0.07099	—	−0.13144[a]
		(0.05360)		(0.07050)
Parent age	—	0.00159	—	0.00205
		(0.00488)		(0.00575)
Parent sex	—	−0.11062	—	0.22247[a]
		(0.10743)		(0.12504)
Constant	—	1.15569**	—	1.35920**
		(0.42989)		(0.52519)
Families:	259	948	187	948
Unequal divisions:	259	259	187	187
Observations:	397	1400	293	1400
Average log-likelihood:	−1.558	−0.967	−1.707	−0.825
R^2/χ^2_{10}	0.010	13.656	0.012	18.896

Notes: The sample is all multichild families for which income tax returns were found for all children (Sample 4). In columns (1) and (2) the exact definition of equal division is used. In columns (3) and (4), equal division is defined to have occurred if all siblings inherit within ± two percent of average. Earnings and parental income are in $100,000's. Asymptotic standard errors are in parentheses. The OLS results include R^2; the χ^2 statistics for a model with zero-slope coefficients are reported for the generalized tobits. The generalized tobit estimates are generated using simulated maximum likelihood with kernel frequency smoothing. The smoothing parameter is 0.10, and 50 random draws are taken for each random variable. Both $(\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})$ and ω_j are normal; σ^2_ω is normalized to 1.0.

[a] Significant at 10 percent.

[b] Children's characteristics are deviations from within-family averages.

* Significant at 5 percent.

** Significant at 1 percent.

that the slope coefficients are zero cannot be rejected at conventional significance levels.

Applying the model with the relaxed definition of unequal division in column (4) provides a better fit; the chi-squared statistic is significant at 5 percent. As before, the generalized tobit estimate of the effect of children's earnings is smaller than found with OLS (in column (3)). In addition, parental income, whether a spouse survives, the number of children, and the parent's sex are significant determinants of psychic costs and consequently the decision to divide an estate unequally. To examine the marginal effects, consider a simulation of the probability of unequal division using 500 draws for each of the ε_{ij} and ω_j. The simulated probability of unequal division for a male decedent who is survived by a spouse is 0.307 (the other parental variables are evaluated at their means, and the intersibling differences are evaluated at the means of their absolute values). The probability falls by 0.034 for a female decedent and rises by 0.048 if the male decedent is not survived by a spouse. A one standard-deviation (just under \$40,000) increase in the mean absolute value of children's earnings differences increases the unequal division probability by 0.015. Although the estimates indicate that the probability of unequal division rises by 0.004 with a \$100,000 increase in parental income, this result is not robust to using the parent's gross estate as the measure of resources (a \$1 million increase in the gross estate lowers the unequal division probability by 0.002). Although the direction of the effect of parental resources on estate division depends upon the resource measure used, it does appear that the magnitude of the resource effect is very small. Moreover, there is essentially no change in the estimate of intersibling compensation under different measures of parental resources.

The generalized tobit estimate of intersibling compensation also is robust to several other sensitivity checks. Specifically, there is little change in the estimate when ω_j is drawn from a lognormal distribution, when the psychic costs in (8) are squared, and when the model is estimated with 200 draws for each ε_{ij} and ω_j. Furthermore, as expected, the bias in the OLS estimate is more severe when two-child families are analyzed in isolation; the fixed-effects estimate of intersibling compensation reported in the previous section is more than twice the magnitude of the generalized tobit estimate (coefficient = −0.144, standard error = 0.021). Relaxing the assumption of symmetric concern by permitting a correlation between ε_{1j} and ω_j produces an inconsequential change in the effect of children's earnings in two-child families. Similarly, investigating several other possible implications of more complicated forms of asymmetric concern does not substantially affect the estimated amount of intersibling compensation.[22] Finally, it is interesting to note that in approximately half of the 259 families where unequal division occurs earning differences between siblings are reinforced by larger bequests to children with higher earnings.

C. *Additional Results*

Altruistic models also predict intergenerational compensation, that is, a negative relationship between the average bequest to children and children's average earnings. Between-family models—equation (5) averaged across the N_j children in each family—are necessary to estimate the magnitude of this effect, which unlike its direction is theoretically indeterminate.[23] Estimates of these models are not necessarily subject to selectivity bias as long as ε_{ij} and ω_j are uncorrelated.[24]

Columns (1) and (2) in Table 5 present estimates of between-family models for families with unequal and equal division, respectively. Several functional forms were estimated (linear, cubic, semi-log, double-log, and a

[22] The OLS estimate of intersibling compensation in a regression derived from a two-child quasi-homothetic parent utility function with asymmetric weighting of children is only slightly higher (−0.375, standard error = 0.205) than was estimated in the fixed-effects model for two-child families. Furthermore, modeling within-family inheritance differences with nonlinear functional forms does not substantially affect the estimated compensation, as it may under more complicated asymmetric utility. Finally, although some parent characteristics do affect within-family bequest differences (as they would be expected to do under asymmetric concern), their inclusion does not alter the effect of children's earnings. Statistically significant parent effects are a surviving spouse and the number of children, both of which reduce within-family bequest differences. Bequest differences are larger if the decedent had a closely held business.

[23] The negative compensatory effect of $\bar{h}_{\cdot j}$ on $\bar{b}^*_{\cdot j}$ is offset by its positive income effect through $\hat{Y}_j$. The magnitude of the latter depends upon functional form and is unknown a priori. However, the net effect of $\bar{h}_{\cdot j}$ is negative if parental consumption is a normal good.

[24] If ε_{ij} and ω_j are uncorrelated then no selection bias can arise from a correlation between $\bar{\varepsilon}_{\cdot j}$ (an error term in the between-family bequest function) and the $(\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})$ terms in (7) because $E[\bar{\varepsilon}_{\cdot j}(\varepsilon_{ij} - \bar{\varepsilon}_{\cdot j})] = 0$. There may be correlation between the unobserved minimum demand component of parental own consumption (η_j) and ω_j, but this would imply a difference in average bequests to children between families experiencing equal and unequal division. However, as previously noted, $b^{**} \approx \bar{b}^*$ on average in the EITM data. Finally, controlling for selection bias in the between-family models produces virtually no change in the results.

TABLE 5—ADDITIONAL INHERITANCE MODELS

Dependent variable	Families with unequal division Average inheritance of children (1)	Families with equal division Average inheritance of children (2)	All matched children Inheritance (3)
Earnings[a]	−1.30674	−0.19616	−0.51273
	(1.03119)	(0.16129)	(0.33295)
Earnings[a] squared	1.67793*	—	—
	(0.65780)		
Earnings[a] cubed	−0.29374**	—	—
	(0.10486)		
Age[a]	−0.00089	0.02605**	—
	(0.01707)	(0.01003)	
Sex[a]	−0.23341	−0.46983*	—
	(0.48539)	(0.23413)	
Spouse[a]	−0.84774	−0.26351	—
	(0.52856)	(0.30390)	
Grandchild[a]	0.08601	0.07070	0.08524
	(0.15191)	(0.09588)	(0.04951)
Parent income	0.14620**	0.15752**	0.21538**
	(0.04579)	(0.02708)	(0.05719)
Surviving spouse	−1.09362**	−1.55441**	−1.56870**
	(0.28302)	(0.16667)	(0.12665)
Number of children	−0.67397**	−0.47628**	−0.63237**
	(0.14494)	(0.08723)	(0.04818)
Constant	5.26712	3.62506	3.49620
	(1.01316)	(0.58840)	(0.50765)
Adjusted R^2:	0.191	0.208	0.159
F statistic:	7.099	23.627	113.734
Sample:	4	4	3
Children:	—	—	4153
Families:	259	688	2020

Notes: Dollar amounts in \$100,000's. Standard errors are in parentheses (corrected for the correlation in error terms among siblings in column (3). Column (1) gives the average bequest to children in each family with unequal division. Child variables are the within-family averages. Estimates are weighted least squares using the number of children per family as weights. Column (2) is the same as column (1) except the bequests are from families with equal division. Column (3) gives inheritances of children. Estimates are two-stage least squares with earnings identified by an age quadratic, sex, and marital status. The regression also includes parent's sex and age (age is significant).

[a] Children's within-family average in columns (1) and (2).

* Significant at 5 percent.

** Significant at 1 percent.

specification mimicking that of Tomes' (1981 Table 4), and the ones in Table 5 were chosen because they yield the strongest evidence of intergenerational compensation. In principle, the use of different functional forms for unequal and equal divisions is not inconsistent with the generalized model because it permits different bequest functions for $\bar{b}_j^*$ and b_j^{**} despite their being derived from the same utility function.

The cubic specification estimated for the average bequest when estate division was unequal suggests that to the extent altruism exists in the data, it is children with lower earnings that experience bequest reductions as their earnings increase.[25] The decedents who made equal bequests to their children reduced these bequests by \$0.196 if children's average earnings were \$1 higher, but this estimate is not

[25] The estimates imply a \$0.60 lower average bequest per each additional dollar of children's average earnings if earnings are below \$45,000, but a \$1.01 higher bequest if earnings are between \$45,000 and \$145,000. I am grateful to a referee who points out that this pattern is consistent with a model containing both altruistic and exchange components, and with the latter dominating as children's earnings increase.

significantly different from zero. Also, recall that the amount of intergenerational compensation is considerably smaller than these point estimates indicate because the models are being estimated in units of annual, not lifetime, earnings.

An alternative interpretation of the lack of strong altruistic evidence in both the within- and between-family models is that compensating behavior is undetectable with the EITM data because of measurement error in parent income and child earnings. However, Monte Carlo experiments suggest that while measurement error can create a large bias toward zero in the estimate of the effect of earnings in within-family models (even with three years of observed earnings), it is unlikely to have led to the present results unless the transitory component of children's earnings is considerably larger than has been measured in other data. The simulation is described in Appendix D (available from the author upon request).

While the transitory component of children's earnings is less of a problem in the between-family analyses because earnings are averaged over several siblings as well as up to three years, measurement error in parent income would generate a positive bias in the coefficient on average child earnings because of the positive correlation between the latter and unobserved parent income. However, simulation exercises again indicate that it is likely this measurement error would not generate the results obtained from the EITM data. Moreover, the bias can be mitigated by the use of instrumental variables as in column (3); note that this specification combines children receiving equal and unequal divisions. Using sex, marital status, and an age quadratic to identify children's earnings yields a negative point estimate on child earnings which is significant at 12 percent, but has little effect on the parent-income coefficient.[26] Alternative measures of parent resources also were considered. Replacing parent income with imputed lifetime income (based on a regression using the Panel Study of Income Dynamics) or parent net wealth, both arguably better measures of the parent's lifetime resources, makes little difference in the effect of child earnings in columns (1), (2), and (3), and in the generalized tobit models of Table 4.[27]

Reconsidering the models in columns (1), (2), and (3) of Table 5, there is some evidence of more intergenerational compensation among decedents who either are survived by a spouse, have gross estates over \$5 million, or who do not create trusts.[28] This last finding is in contrast to the weak evidence that those not leaving trusts reinforce intersibling differences. Finally, estimating the specification used in column (3) for one-child families alone provides the largest point estimate of the degree of intergenerational compensation (−1.136, standard error = 1.436) obtained in the paper.

[26] The OLS estimate is +0.156 (standard error = 0.155). Issue may be taken with the instruments selected to identify child earnings because the excluded variables may directly affect the parent's bequest decision. Unfortunately, these are the most reasonable instruments available in the data. These results are sensitive to the exclusion of parent's age: omitting it reduces the estimate of intergenerational compensation to −0.067 (standard error = 0.320). In contrast, the compensation estimated in columns (1) and (2) is insensitive to the inclusion of parent age.

[27] However, the coefficients on imputed lifetime income are 41.4 to 173 percent larger than those on parent income presented in columns (1), (2), and (3), of Table 5. The imputation is based on 22 years of income data for nonfarm parents over age 40 in the PSID's Survey Research Center cross section (n = 747). Their lifetime income (discounted present value of annual income divided by 22) is regressed on variables which have a corresponding measure in the EITM data, most importantly their average income over the last two years, their 1989 net wealth, and the earnings of their children (averaged over all children remaining in the PSID and over the last three years). The coefficients on these variables are 0.347 (standard error = 0.016), 0.008 (0.002), and 0.227 (0.042), respectively. The regression also includes a quartic age profile, sex, whether married, the number of children, and whether ever self-employed. The adjusted R^2 is 0.673.

[28] The evidence of greater compensation of child earnings occurs when a parent survives [−0.392 (standard error = 0.187) and −0.763 (0.436) in Table 5, columns (2) and (3), respectively] and when no trusts are created [−0.292 (0.201) and −0.850 (0.441) in columns (2) and (3), respectively]. Excluding gross estates greater than \$5 million reduces the estimated effect to −0.090 (0.139) and −0.254 (0.279) in columns (2) and (3), respectively. Excluding decedents who made gifts prior to death does not produce stronger evidence of intergenerational compensation.

Although not precisely estimated, this larger estimate is consistent with the theory that the strength of altruistic motives relative to exchange considerations increases when parents are unable to credibly threaten children with the possibility of bequeathing less to them in favor of their siblings (Bernheim et al., 1985).

V. Conclusions

Federal estate tax returns describe the bequest patterns of the top wealth holders who control a substantial portion of U.S. wealth. Inheritance data from these returns matched to income tax return data for child beneficiaries provide little evidence that bequests are compensatory. First, the majority of all wealthy decedents bequeath equally to their children. This result clarifies the estate division debate and motivates a theoretical generalization of the standard altruistic model. Second, although large earnings differentials between siblings make an unequal estate division more likely, and when an estate is unequally divided inheritances do provide some compensation to children with low earnings, both effects are very small. Third, lower average earnings of children generally do not induce statistically significant higher bequests from their parents. However, there is somewhat more evidence of altruistic bequests among one-child families, exactly those families for whom exchange motivations are likely to be most weak.

It is important to note that the evidence suggesting that the altruistic model does not represent the bequest behavior of most wealthy families comes from a data set which excludes those expected to be nonaltruists. Decedents breaking the presumed altruistic connection, either because of being childless or by being at a corner solution in which they desire to leave a zero bequest, were excluded. Capital-market imperfections which may constrain the decisions of the poor are not constraints on the decisions of the decedents in this sample. At the same time, it is fitting to remember that despite the advantages of the EITM data for examining intergenerational transfers, they do contain several limitations. Recall that the data refer to natural born and adopted children; unfortunately, stepchildren are not included. Also, the earnings data are subject to measurement error, disinherited children are not necessarily reported, wealth passed through trusts cannot be traced to its new owners, and only the last three years of inter-vivos gifts are known. Although I have argued that these factors have limited consequences, it is important to keep these issues in mind when interpreting the results.

The results have several implications. The findings add to the evidence that there is a tenuous empirical micro-foundation for the behavioral model upon which the Ricardian-equivalence propositions rest. Also, government efforts to achieve intergenerational redistribution will not be neutralized by bequests. Finally, there is little behavioral basis for the theoretical argument that, because inheritances compensate children with low earnings, estate taxation increases inequality (for example, Becker and Tomes, 1979).

Like the available evidence on inter-vivos gifts, which also has not revealed strong compensatory behavior, these results are consistent with an exchange model of intergenerational transfers, especially in light of the results for one-child families. However, the results also are consistent with altruism that is either complicated by the strategic behavior of parents and children or expressed in the context of other parental goals. For instance, parents may desire to equalize the opportunity sets (bequests, inter-vivos gifts, and human-capital investment) of their children. Subsequently, the decisions of the children lead to unequal outcomes, but these are not routinely redressed by parents unless they are perceived to arise because of events beyond the children's control. As an example, parents may not normally compensate a child with low earnings, but would if he became unemployed.[29] In addition, the results do not rule out the possibility that parents make altruistic transfers in response to a child's specific needs which occur at the purchase of a first house, the birth of children, or other points in the life cycle.

[29] Consistent with this possibility is Robert F. Schoeni's (1991) evidence that inter-vivos transfers decline if the child receives unemployment compensation. However, most of the evidence presented by Hayashi et al. (1991) rejects risk sharing within the family.

These issues, as well as those implicitly raised by the qualifications to the present results, provide directions for further research. Of particular importance, data on inter-vivos gifts received by heirs are necessary to concretely determine whether earlier gifts reduce the prevalence of compensatory bequests. Also, such data could be used to develop an understanding of the entire life-cycle profile of intergenerational transfers, and would contain the additional earnings information essential to reduce measurement error bias.

Finally, because the economic and demographic characteristics of siblings are not identical, any theory in which bequest behavior is conditioned on such characteristics necessarily runs aground on the empirical observation of equal division. This includes all of the recent theories based on altruism or exchange. Indeed, the only parent preferences consistent with equal division are those in which the bequest is "accidental," or those in which utility is gained from the size of bequest. However, the former theory offers no refutable hypotheses concerning estate division, and the latter theory predicts no unequal division, which is inconsistent with the observation that a substantial minority do bequeath unequal amounts to their children. Thus, theoretical explanations of estate division practices remain to be found.

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