CHAPTER 10. EFFECTS OF CHANGING AGE STRUCTURE AND INTERGENERATIONAL TRANSFERS ON PATTERNS OF CONSUMPTION AND SAVING

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1. Introduction

This study examines the issue of age-structure transitions and intergenerational transfers from a macro rather than a micro perspective: to what extent might changing age structure, brought about by the demographic transition and subsequent baby booms and busts, have contributed to economic fluctuations experienced in countries around the globe in the last century? And to what extent might intergenerational and interhousehold transfers have confounded attempts to quantify this relationship? These issues are explored here using data for the United States throughout the 20th century, as the United States moved from "developing" to "developed" status.

The post World War II baby boom—and then bust—in many Western nations, as illustrated in Figure 1, provides a natural experiment for examining the effects of a significant population "bulge" as it moves through the life-cycle. A simple life-cycle model suggests that there is a marked age-related fluctuation in the proportions of income consumed and saved over the life-cycle, as illustrated in Figure 2, which when superimposed on Figure 1 would suggest major shifts in patterns of expenditure and savings. Individuals are thought to overspend relative to their incomes (dis-save) at younger ages and again in retirement, with a period of saving during the prime years. If such a pattern exists at the micro level, does it carry through when the data are aggregated over individuals, and if so how significant is it in affecting macro level economic variables?

Similar issues have been addressed in a wide range of studies focused on adults aged 15 or 20 and older, like those of Angus Deaton and Christina Paxson. (See, for example, Deaton and Paxson 1997.) But another strand of the literature on these types of age structure effects has focused on the impact of children on savings rates and economic growth in less developed countries (LDC): what is referred to as the "dependency effect" (Leff 1969; Mason 1981, 1988; Fry and Mason 1982; Collins 1991; Kelley and Schmidt 1995). There are no clear answers there, however: it appears that the literature has highlighted two puzzles in recent years, both of them apparent reversals in the sign of the impact. The first is a very well documented reversal in the estimated effect of youth dependency on savings and

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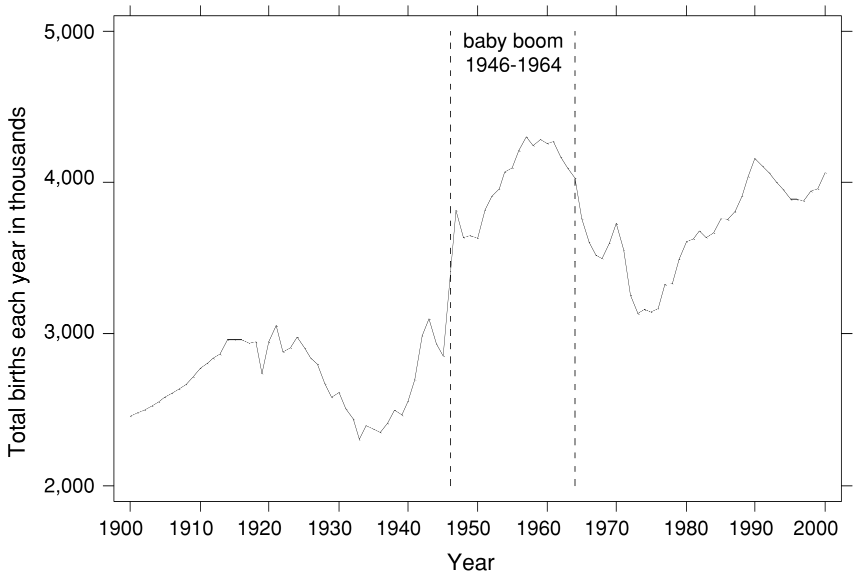

Figure 1. Pattern of births in the United States in the 20th century.

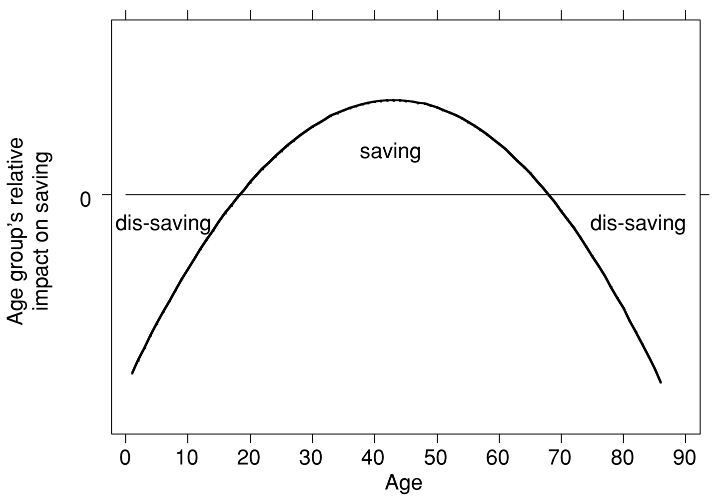

Figure 2. The hypothesized "life-cycle" pattern of consumption and saving, as a share income.

economic growth worldwide: what was a negligible or even benign effect in the 1960s and 1970s seems to have turned negative in the 1980s (Kelley and Schmidt 1994). The second is an apparent variation in the sign of the effect, in moving from LDCs to DCs (developed countries). That is, while these studies in general support the idea of a significant negative (short-term) effect of high birth rates on savings rates and economic growth rates in lower income countries, they suggest that this negative effect may be ameliorated as income levels rise.[1] Does this mean that little can be learned from the study of dependency rates in LDCs, to assist in estimating age structure effects in Western nations, and vice versa?

Not necessarily. Given the right data set it is possible to control for—and even quantify—age-specific effects that change with level of development, although this has not, to the author's knowledge, been done in any studies to date. Inconsistent estimates of the dependency effect may be due to inadequate measures of population age structure. Until very recently the standard measure in the dependency literature was simply the population share aged 0–14—or that share relative to the share aged 15–64—and there seems reason to suspect that measure might be capturing something other than, or in addition to, youth dependency. Why? Because the sizes of various age groups tend to be highly correlated in any population.[2] Parallel movements in multiple age groups make it impossible to determine if low savings rates are actually caused by the dependents themselves, or by adult age groups with whom their size is correlated, without adequate controls for the relative sizes of those adult age groups.

In addition, recent research suggests that even the 0–14 age group cannot be treated as homogeneous, although this is a standard approach in the dependency literature. The ratio of expenditure to income is significantly affected by the age of children in a family, even holding constant the age and real income of adults in that family (Lazear and Michael 1988). Holding income and other factors constant, parents in the United States are found to spend about 15 percent more on teens as opposed to younger children; this result is supported by data reported in Lino (1998). Thus, in addition to controlling for many heterogeneous adult age groups in dependency studies, it may be necessary to recognize not just the presence of children, but also their age distribution.

None of this is to say that there is no negative youth dependency effect. The point is that it seems doubtful that we can even know if there is any effect at all, much less its direction, until or unless we can control for more than just one or two age groups. Fair and Dominguez (1991) introduced a unique way of doing just that, and it seems pertinent that two recent studies making use of their technique did not seem to find "inconsistent" results between LDCs and DCs, but rather variations in effects as children aged. Higgins and Williamson (1997) examine what they term "population dynamics" using the full age distribution, examining the effects of a young population as it passes from childhood to working age. In doing so they identified what appear to be very significant economic effects: an initial period of what they term a "demographic burden" (when the new population is very young) followed by a "demographic gift" as the additions to the population age into productive

[1] Kelley and Schmidt (1995:544), in their extensive analyses of 99 DCs and LDCs between 1960 and 1990, found that the *overall net effect* of population growth on economic growth sometimes appears to move from negative to positive as per capita income rises—at least prior to the 1980s. They used current and lagged birth and death rates to attempt to capture dynamic age structure effects that they felt were missed by more static equilibrium analyses based on more traditional measures of dependency, like the share of population aged 0–14. But even some of the more traditional analyses seemed to find differential effects, for example Collins (1991); Taylor (1995); and Taylor and Williamson (1994).

[2] Examples of such correlations in the U.S. population, and their changes over time are presented in Macunovich (2001).

maturity. As in Higgins (1998), an initial period of rapid increase in investment seems to be generated by population growth, creating a substantial negative current account balance, but this is followed by a period of strong labour force growth accompanied by an increase in domestic savings and a transition to positive current account balances. These findings are consistent with Barlow (1994) and Kelley and Schmidt (1994, 1995), who find that long- and short-term effects of population growth differ, moving from negative to positive as a population bulge ages.

The purpose of the present analysis is to identify any effects of changing age structure on patterns of consumption and savings, while controlling for the sizes of all age groups in the population using the Fair and Dominguez technique, and permitting the age-specific effects to vary by level of development (per capita income). This has not been done in previous studies, possibly because in cross-national comparisons variations in per capita income tend to be confounded with a host of institutional and cultural differences. This study attempts to sidestep that problem by using cross sections of U.S. state-level data throughout the 20th century, tracing the United States from its "less" to its "more" developed status in the personal consumption expenditures (PCE) of up to 51 states (including the District of Columbia) with a greater degree of internal consistency in culture and institutions, than can be found in international cross sections.

In addition, the analysis presented here incorporates another aspect of changing age structure which has been omitted in all previous analyses: the potential effects of age structure—"relative cohort size"—on individuals' income relative to their material aspirations—their "relative income." These effects were first hypothesized by Richard Easterlin (1987), and are discussed in more detail in Section 4.

What this analysis finds is a more complex pattern of age structure effects than expected, a pattern that does indeed appear to vary both by level of income and by relative cohort size. The estimated effects are highly significant, suggesting that changes in age structure have caused PCE in the United States to vary by about 25 percent during the 20th century, holding constant all other factors including population size, real income, and its growth rate.

The next section addresses problems inherent in both micro and macro level analysis, while Section 3 explains the potential significance of intergenerational transfers and Section 4 explains the relative cohort size concept. Section 5 describes the data and method used in this study; Section 6 demonstrates the effect of using these data in a few model formulations from previous analyses in the dependency literature; and Section 7 presents results from applying a new method and model to the U.S. state-level data. The chapter concludes in Sections 8 and 9 with an application of these predicted age structure effects in a simulation of consumption behaviour in the United States given the population changes that occurred throughout the 20th century.

2. Micro versus Macro

It must be emphasized at the outset that any age structure effects identified at the aggregate level cannot be interpreted as support for, or refutation of, the standard life-cycle model of saving and consumption. We should neither assume nor expect to see the standard "hump-shaped" pattern of saving and spending in the aggregate (as emphasized in Higgins 1998, p. 351). That is, for example, the longevity of elderly parents—and therefore their continued presence in the age structure—might induce middle-aged adults to spend more in supporting them than they otherwise would have, or alternatively might induce those middle-aged adults to *save* more in the expectation of their own increased longevity. The presence of children in an extended family might induce additional saving, or spending, on the part of aunts, uncles, and grandparents. In this sense, the life-cycle model is simply a starting point generating the possibility of age structure effects, rather than the model being tested here. And conversely, this implies that any coefficient estimates of life-cycle spending/saving patterns based on individual household micro data cannot simply be aggregated to estimate macro level effects—although this has been a common practice in the savings literature for industrialized nations.[3]

Weil (1994) addresses this aggregation issue in explaining the differences between measured age patterns of savings at the micro and macro levels, that he feels result from intergenerational responses to the bequest motive. He posits that increased saving by elderly households planning bequests, as measured in micro level surveys, is masked at the macro level because of reduced saving by adult children expecting to receive those bequests. He concludes that "... one cannot use the mean saving of people at different ages (or any other coefficients that come from micro data that do not account for members of other generations) to forecast changes in the aggregate saving rate in response to changes in the age structure of the population" (p. 67).

Weil's findings in support of that hypothesis argue strongly for the use of macro level data to estimate effects of changing population age structure. Another study that supports this type of intergenerational effect is Attanasio (1998), where the savings profile of cohorts born between 1920 and 1939 appears to be "shifted down" relative to that of preceding and subsequent cohorts. Attanasio suggests that this shift might have resulted from baby boom-induced intergenerational transfers. In addition, members of the 1920–1939 birth cohort may have reduced their savings because as the "sandwich generation" they found themselves simultaneously caring for both children and ageing parents who may not reside in the same household.[4] The use of micro level data like the Consumer Expenditure Survey (CES) will miss much of this type of intergenerational effect, to the extent that it occurs across rather than within households.

But as Taylor and Williamson (1994) point out, macro data present problems as well: "The use of multi-country cross sections for short periods (or even a single year) raises the possibility of omitted variable bias... The use of long time series for a few countries (or only one country) raises the question of the structural stability of the savings equation and inclines a model to track poorly over the short to medium term." (1994:360 footnote 14). The state level cross sections used in the present analysis lie midway between micro and macro, capturing the inter-household effects otherwise available only in national macro

[3] See, for example, Auerbach, Cai, and Kotlikoff (1990) and Bosworth, Burtless, and Sabelhaus (1991).

[4] I thank my former colleague at Williams College, Roger Bolton, for this insight.

data, but using aggregations exhibiting more cultural and institutional similarity than can be found in international cross sections.

3. Inter-Household Transfers and Their Role

Intergenerational—and therefore usually interhousehold—transfers have been documented to occur among all income groups and to be particularly significant from older to younger generations. Sumon Baumik, in another chapter of this book, emphasizes the significance of such transfers, and states that "an event like child birth can induce transfers equal to about 30 percent of the average annual income of the recipient's household." (p. 112) Other examples of such interhousehold transfers, as pointed out by Baumik, might be gift giving at graduation or marriage. Even the expenditures of family and friends in dressing for and travelling to such functions would be expenditure induced by the age of the younger group, rather than by the ages of persons attending.

A childless adult might choose to assist financially a younger sibling with "start-up" costs for the younger sibling's children, or might set up a savings account in the children's names in anticipation of future educational expenses. Similarly parents might draw down on their savings in order to provide their adult children (no longer living at home) with cash for a car or a deposit on a house. This would be behaviour induced by the age and presence of the children, rather than by the age of the parents or other relatives; but the children's presence would not be detected in micro level surveys of the expenditures of individuals living outside the children's own households.

If, as suggested by Welch (1979) and Macunovich (1999), the incomes of young adults in large birth cohorts are adversely affected by cohort size, while at the same time the incomes of their parents are favourably affected, such cross-cohort giving might produce patterns of savings and consumption not predicted by models dealing with households in isolation—patterns related to the age distribution in the *total population*, rather than to the age of a household's own head. Similarly, when adult children contribute to the nursing home care of their elderly parents, that expenditure's relationship to the parents' age group would be missed in a household-level survey, but picked up in an analysis at a more aggregate level. Even expenditures of time on elderly parents—to the extent that they reduce paid work hours among the caregivers—could show up at the macro level as a change in the share of consumption out of income.

Why might these effects—and especially the effect of children—vary between lower and higher income countries? Given imperfect capital markets and an absence of government provisions for old age security in early stages of development, children themselves are effectively a form of saving on the part of their parents—their "old age security". Thus, it is perhaps to be expected in developing countries that high youth dependency rates will have an immediate negative effect on formal savings rates. But the effect of children in more developed economies is open to question. As material aspirations increase along with economic development, and parents begin to opt for "quality" over "quantity" in children, higher levels of education become mandatory, and parents begin to have more ambitious plans for their children as young adults. (See, for example, Kelley and Schmidt 1994: 24.)

It seems possible, then, that with increasing levels of development parents might be induced to save more, rather than less, when they have children, in anticipation of higher costs in the children's late teens, and when they set up on their own as young adults. Grandparents, as well, may save more, either to assist with the costs of higher education, or through an increased bequest motive as suggested by Collins (1991). Even college endowments are a form of saving by society in general, induced by the presence of children who will need education in the future. In addition, economists at least as far back as Malthus have recognized that children cause parents to work harder,[5] so that although absolute consumption expenditures may rise in the presence of children, there may be little increase in consumption relative to income. The net effect of children on aggregate patterns of consumption and savings is thus an empirical question.

4. An Additional Wrinkle: Relative Cohort Size Effects

There is an additional complication that arises in analyses of this type, because of changing age structure's potential to exert not only simple compositional effects on consumption (through life-cycle changes in behaviour), but also to be responsible for *changes in age-specific effects over time.* This phenomenon has not been addressed in the savings and consumption literature to date. It arises from the impact of *changing relative cohort size on young adults' relative income*, as first pointed out by Richard Easterlin (1987), and later substantiated in the United States by Welch (1979); Berger (1984, 1985, 1989); and Macunovich (1999), among others. Similar effects have been documented in other countries, both developed (Korenman and Neumark 2000) and developing (Higgins and Williamson 1999).[6]

The phrase "relative cohort size" refers to the size of a birth cohort of young adults relative to the size of their parents' birth cohort. "Relative income," in turn, refers to the earning potential of young adults relative to their material aspirations, which Easterlin hypothesized would be a function of the standard of living experienced while at home with their parents. The mechanism of transmission from relative cohort size to relative income is *imperfect substitutability in the labour market*. That is, increases in the size of entering cohorts in the labour market might be expected—through standard demand–supply effects—to reduce wages, but this effect does not occur uniformly throughout all experience groups. Because young inexperienced workers are very poor substitutes for older experienced workers, the increase in their own supply will depress their own wages, but may even increase older workers' wages. And because the older workers are, in the aggregate, the younger workers' own parents, this depresses the young workers' own relative wages and hence their earning potential relative to their material aspirations—their relative income, and thus the average relative income of all young adults at the aggregate level.

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[5] Malthus wrote in his 1817 appendix to *An Essay on the Principle of Population*: "If it were possible for each married couple to limit by a wish the number of their children, there is certainly reason to fear that the indolence of the human race would be very greatly increased . . . " (Page 369 in the Cambridge University Press 1992 edition selected and introduced by Donald Winch.)

[6] I am indebted to members of this IUSSP workshop—especially to Nancy Folbre, for suggesting explicit controls for relative cohort size effects.

Easterlin hypothesized that this reduction in relative income would cause compensatory changes in age-specific behaviour in larger birth cohorts, adjustments such as reduced marriage and fertility and increased female labour-force participation. These changes in age-specific behaviour would in turn very likely affect patterns of age-specific consumption relative to income, as young people struggle to maintain their desired standard of living, and parents make transfers to young adult children to supplement their reduced wages. To the extent that this is the case, any study of age structure effects on consumption must control for changes in relative cohort size, as well as in the overall age structure.

Although these relative cohort size effects are difficult to study at the micro level, due to the absence of adequate income data describing the parents of both partners in a marriage, and the impact of other factors on material aspirations, they are much more straightforward in national- and state-level data, where it is not necessary to link specific parents with specific children. If the effect is there, it should emerge at the generational level when changes in the average income and cohort size of, say, prime age workers, are compared with changes in the average income and cohort size of young adults.

5. Data and Methods

This study addresses the issue of age-related patterns of consumption in a new way, using state-level cross sections of PCE developed for the United States by Lebergott (1996). For five dates in this century—1900, 1929, 1970, 1977, and 1982—he provides data that have three significant advantages over other data used in the past to analyse age structure effects on macroeconomic measures. First, they permit the analysis of 48 to 51 (including Alaska, Hawaii, and the District of Columbia) "sufficiently similar" areas as suggested by Taylor and Williamson (1994), in order to minimize bias from omitted variables. Second, they permit the examination of the United States over a long time span but with less danger of bias due to the autocorrelation that occurs in annual time-series data. Stoker (1986) has demonstrated that the most common method of dealing with autocorrelation—the inclusion of leads and lags of the dependent variable—produces misleading estimates because the leads and lags normally include information on the only slowly changing patterns of age structure, and thus tend to reduce the estimated significance of age structure variables. And thirdly, these state cross sections permit comparisons between recent U.S. experience as an industrialized nation coping with baby booms and busts, and the United States at the beginning of the 20th century when it was itself virtually a "developing economy" still experiencing its demographic transition.

Lebergott developed several hundred new series for components of PCE annually for the years 1900–1929—and at the state level for 5 years—that are directly comparable with official Bureau of Economic Analysis (BEA) series as revised in 1993, in both current and constant dollars. These data prepared by Lebergott are far more comprehensive than data provided by the CES. As he points out the latter are based on interviews with "less than one-thousandth of one percent of American 'consumer units'" in which "individual members of households try to remember expenditures in the prior year" (p. 130). The 1984 survey, for example, "understated U.S. food and clothing expenditures by $173 billion. Not to mention $33 billion for house furnishings, $28 billion for alcohol, and $46 billion for entertainment (pp. 129–130)," while the census rent sample "was one thousand times greater than that of the [Bureau of Labour Statistics] BLS" (p. 131).

Lebergott's data began with BEA national income account totals that were then allocated to states. It is important to note that in no case were his allocation methods based on age distributions within the population. Rather, they were derived from census data on production and expenditures, as well as (for 1900) distributions of workers by occupation and service income. Lebergott's data are described in more detail in the Appendix, and summary statistics are presented in Table 1.

These state and national expenditure figures are supplemented with detailed population data for states in each of these years, and for the United States in the entire century, provided by the Bureau of the Census. The analysis makes use of detailed age breakdowns of the total population, rather than simple dependency rates or birth rates, following the lead of Fair and Dominguez (1991), and recent analyses by Higgins and Williamson (1997), and Higgins (1998).[7] This minimizes problems associated with identifying appropriate age groupings. For example, including the 25–34 age group with assumed savers in the 35–44 age group would dilute the measured effect of that older group, if in fact the 25–34 year olds were dis-savers rather than savers.

Ideally a large number of age groups would be separately identified as independent variables in the analysis to avoid this age group identification problem. But a model that includes a large number of age groups in order to overcome the possibility of erroneous groupings, encounters a problem of severe multicollinearity that calls into question the accuracy of any individual coefficient estimates. And problems of multicollinearity are compounded by the marked loss of degrees of freedom in estimating those coefficients, as the number of age groups is increased: an important consideration in time series analyses. As observed by David (1962), "Age varies continuously and there are few convenient demarcations between age groups with significantly different behaviour patterns."

In order to circumvent these associated problems of age group identification and multicollinearity, the present analysis uses the Fair-Dominguez (1991) method of parameterizing multiple age groups, which is in turn based on Almon's (1965) distributed lag technique. This methodology has subsequently been adopted in more recent studies, as mentioned earlier, and is described in more detail in the Appendix. In general terms, it is one that permits the estimation of coefficients on single year population age shares by constraining those coefficients to lie along a polynomial. The coefficients φ_j on J population age shares p_j are assumed to enter the consumption equation in the form

$$\sum_{j=1}^{J} \varphi_j p_j \tag{1}$$

[7] It must be emphasized, in response to an earlier reviewer's comments, that the inclusion of the full age distribution in a more aggregated consumption/savings model does not in any way imply or necessitate decision-making on the part of children themselves, with regard to their patterns of consumption. It simply allows for the fact that children in one household might affect the spending patterns of individuals in other age groups and/or households.

Table 1. Summary statistics on state-level variables used in the analysis.

Variable (by Year)	Obs	Mean	Std. Dev.	Min	Max
Personal consumption expenditures as a share of total personal income					
1900	48	1.09	0.19	0.76	1.77
1929	48	0.91	0.07	0.73	1.06
1970	51	0.82	0.07	0.66	1.04
1977	51	0.84	0.08	0.63	1.17
1982	51	0.79	0.09	0.61	1.20
Real personal income (in millions of 1987 dollars)					
1900	48	4849.85	6608.08	257.88	35618.23
1929	48	12563.50	17961.98	558.05	99954.68
1970	51	43283.79	52840.65	3370.37	247811.91
1977	51	54474.04	62727.56	5117.60	317822.91
1982	51	62932.26	74108.20	6346.59	395150.09
Total population (in thousands)					
1900	51	1486.67	1539.29	42.33	7268.89
1929	49	2505.61	2519.98	91.06	12588.06
1970	51	3985.32	4314.83	302.34	19964.68
1977	51	4309.02	4559.73	403.44	22352.41
1982	51	4542.44	4846.96	449.61	24820.02
Growth rate of real GDP (percent)					
1900	47	3.98	3.14	−3.76	15.58
1929	49	0.72	1.69	−2.44	3.65
1970	51	3.71	1.35	0.99	7.46
1977	51	2.13	1.97	−4.43	10.05
1982	51	1.10	1.48	−1.92	4.45
Percent urban					
1900	49	33.90	23.54	6.20	100.00
1929	49	46.92	21.01	16.60	100.00
1970	51	66.44	15.14	32.20	100.00
1977	51	67.02	14.99	33.00	100.00
1982	51	67.60	15.00	33.80	100.00
Percent foreign-born					
1900	51	14.83	11.77	0.20	58.90
1929	49	8.97	7.41	0.28	25.35
1970	51	3.44	2.81	0.40	11.60
1977	51	4.34	3.56	0.90	15.10
1982	51	4.34	3.56	0.90	15.10
Z1 (using logged population shares)					
1900	49	−2174.93	370.86	−3047.18	−1320.76
1929	49	−1818.45	353.85	−2720.06	−1200.02
1970	51	−1177.47	272.80	−2560.16	−766.52
1977	51	−1209.02	276.60	−2623.17	−695.44
1982	51	−997.92	254.79	−2256.73	−492.93

Table 1. (*Continued)*

Variable (by Year)	Obs	Mean	Std. Dev.	Min	Max
U.S. averages:					
1900–1925	26	−1991.70	78.02	−2108.43	−1862.23
1926–1950	25	−1534.28	174.02	−1842.22	−1298.83
1951–1975	25	−1192.50	68.89	−1289.75	−1050.85
1976–1999	24	−877.30	84.41	−1030.34	−747.03
		Z2 (using logged population shares):			
1900	49	−205414.50	34096.00	−297895.81	−125640.50
1929	49	−175913.00	33640.09	−268351.50	−116720.30
1970	51	−109927.40	26520.66	−244218.50	−71987.44
1977	51	−115733.20	26706.89	−253996.20	−69216.45
1982	51	−96052.79	24458.84	−220786.41	−49608.36
U.S. averages:					
1900–1925	26	−189452.30	6344.23	−198650.50	−178894.70
1926–1950	25	−150154.80	15745.81	−177205.00	−126569.20
1951–1975	25	−112941.00	7332.17	−125312.90	−99988.94
1976–1999	24	−85372.21	7345.82	−98363.23	−74111.45

which is estimated as a polynomial

$$\sum_{j=1}^{J} \varphi_j p_j = \zeta_1 Z_1 + \zeta_2 Z_2 + \cdots + \zeta_n Z_n \tag{2}$$

in which n is the degree of the polynomial and Z_n is a weighted sum of individual population shares, defined as

$$Z_n = \sum_{j=1}^{J} p_j j^n - 1/J \sum_{j=1}^{J} p_j \sum_{j=1}^{J} j^n \tag{3}$$

Estimating the degree of that polynomial, n, appears in earlier studies to have been based largely on theory, assuming a quadratic at the aggregate level based on the hypothesized life-cycle "hump-shaped" pattern at the micro level. But as emphasized earlier in this chapter, because of interhousehold effects a different pattern may emerge at the aggregate level, and there is more danger in over- than in underestimating the degree of the polynomial. Judge et al. (1985:359–360) state that when fitting an Almon lag, overestimates of the true degree of the polynomial produce estimators that are unbiased although inefficient, whereas underestimates of the true degree produce estimates that are "always biased." For this reason they suggest starting with a higher n, than is assumed to apply in the true model, and stepping down, testing each additional restriction and finally accepting the level that "produces *the last acceptable hypothesis* [their italics]." That procedure has been adopted in the present study, and is documented in results not presented here, but available on request.

6. Reproducing Earlier Results, Using U.S. State-Level Data

The approach adopted in this analysis was first to attempt to reproduce (using U.S. 20th century data) the results of analyses of age structure effects that included countries at different income levels and stages of development, and then to expand on those models by including all age groups in unconstrained versions of the models (i.e., permitting age-specific effects to vary by income level and relative cohort size).

The studies addressed in this way were by Collins (1991); Taylor and Williamson (1994); Taylor (1995); Leff (1969); Fry and Mason (1982), and recent studies by Higgins and Williamson (1997); and Higgins (1998).[8] All of these analyses estimated similar models, except for the fact that the last two tested for effects of the full population age structure rather than just selected age groupings. None permitted the dependency effect to vary by level of income or relative cohort size, although some of the later ones, based on Fry and Mason's (1982) "variable rate-of-growth" model, tested for effects of the growth rate of GDP. They hypothesized that "[t]he dependency ratio [defined as the population aged 0–14 relative to the population aged 15–64] exerts both level and timing effects: a higher dependency ratio reduces the aggregate saving rate, the magnitude of the effect increasing with the rate of growth in real GNP" (p. 427). Fry and Mason's analysis was in turn an extension of earlier work by Leff (1969), who hypothesized (consistent with the analysis in Tobin 1967) that real GNP growth would increase the incomes of younger (saving) relative to older (dis-saving) households, thus increasing overall levels of saving. Thus, the Fry and Mason model of national saving, tested using panel data from seven Asian developing countries between 1962 and 1972,[9] included not just dependency and growth rates, but also their interaction, which was found to have the expected negative effect:[10]

$$
\begin{aligned}
\ln[1/(1-s)] &= -\ln(c) \\
&= \underset{(0.066)}{0.147} - \underset{(1.841)}{3.769\,D} - \underset{(0.130)}{0.526\,f} + \underset{(0.410)}{1.644\,g} - \underset{(18.224)}{69.457\,Dg} + \underset{(1.159)}{4.822\,rg} \\
&\quad + \underset{(0.088)}{0.249}\ \ln[1/(1-s)]_{t-1}
\end{aligned}
\tag{4}
$$

where s is the aggregate rate of saving, the log-odds savings rate was approximated using the negative of the logged rate of consumption c, D is the population under age 15 divided by population aged 15–64, f is the foreign saving rate, g is the rate of growth in real GNP, r

---

[8] The Kelley and Schmidt (1994, 1995) studies are not included in this list because of their use of birth rates rather than specific measures of age structure, to explain effects on growth and savings.

[9] Fry and Mason's sample began with 10 countries, but dropped Pakistan, Sri Lanka, and Thailand because of binding foreign exchange constraints, ending with only seven: Burma, India, Korea, Malaysia, Philippines, Singapore, and Taiwan.

[10] Fry and Mason's full initial model was

$$
\begin{aligned}
\ln[1/(1-s)] &= -\ln[c] \\
&= \alpha_0 + \alpha_1 D + \alpha_2 r + \alpha_3 f + \alpha_4 g + \alpha_5 Dg + \alpha_7 rg + \alpha_8 fg + \alpha_6 \ln[1/(1-s)]_{t-1}
\end{aligned}
\tag{4a}
$$

but they dropped r and fg after finding their estimated coefficients were not significant.

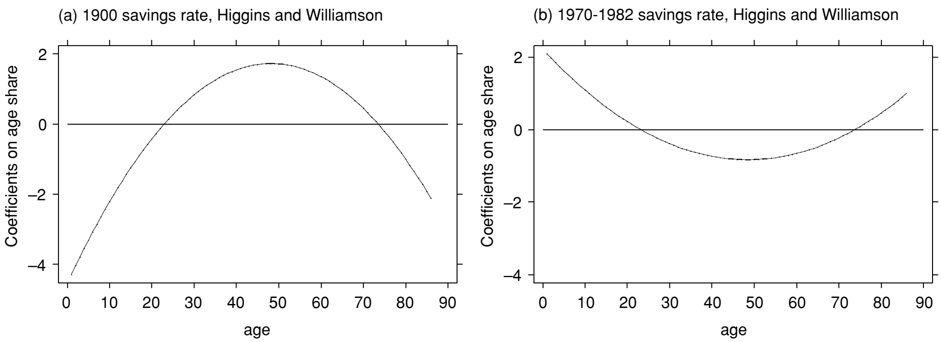

Figure 3. Estimating the Higgins and Williamson (1997) model presented in Table 2 using U.S. state-level data for 1900 and for 1970–1982, illustrating the apparent reversal that occurs in moving from low- to high-income economies.

is the real rate of return on financial assets and the adjusted R^2 was 0.87 (with $t-$statistics in parentheses).

It should be noted, however, that later models based on Fry and Mason have tended to find positive, rather than negative, effects of the interaction term between dependency and the growth rate. These models are described in more detail, and presented with their results, in Macunovich (2001). In addition, Macunovich (2001) presents detailed results from reestimating each of these models using Lebergott's state-level data for the United States at the five available dates (1900, 1929, 1970, 1977, and 1982) in place of the various groupings of LDCs and MDCs (more developed countries) used in the earlier studies. A common pattern that emerges in all of the reestimates is what appears to be a structural break between the earlier and later years, with coefficients supporting the dependency hypothesis in 1900, and "inconsistent" in more recent years. And contrary to Fry and Mason's results, in all of those reestimates the level and rate-of-growth effects of dependency tend to move in opposite directions.

This is perhaps best illustrated graphically, as in Figure 3 and Table 2, using the Higgins and Williamson (1997) model. A negative effect of children on savings in 1900 is demonstrated by the strong negative coefficients that were estimated for the lowest ages in Figure 3a (which are a result of the positive sign on the coefficient estimated for Z_1 in column 3 of Table 2). The hump shape in Figure 3a is the shape expected under the standard life-cycle model—higher levels of consumption out of income, and thus a tendency to "dis-save" at younger and older ages, with lower levels of consumption relative to income in the middle-age groups, the "saving" years. But those expectations are not met in Figure 3b, as was typical of all of the model re-estimates: younger children are estimated here to have a positive effect on savings in more recent years.[11]

[11] The age share coefficients presented in Figure 3b are based on column 6 in Table 2, but the pattern is virtually unaffected by a change to a fixed effects model (as in column 4) or when year dummies are excluded (as in column 7).

Table 2. Original Higgins and Williamson (1997) model estimates (column 2) and reestimates of their model using Legergott state-level U.S. data for 1900 and for 1970–1982, illustrating the apparent reversal in the effects of age structure in moving from the early to late 20th century U.S.—and by implication, from lower to higher income economies.

	Higgins and Williamson fixed effects results	Results using Lebergott U.S. state-level data				
		1900	1970–1982			
			Fixed effects		Random effects	
Z1	0.703**	0.2597***	−0.1217*	−0.1138**	−0.1258***	−0.0770**
		(0.0584)	(0.0687)	(0.0448)	(0.0392)	(0.0372)
Z2	−0.046**	−0.0027***	0.0012	0.0015***	0.0013***	0.0010**
		(0.0007)	(0.0008)	(0.0005)	(0.0004)	(0.0004)
Z1 growth rate*	na†	−0.0219*	0.0232**	0.0376***	0.0239**	0.0232**
		(0.0109)	(0.0104)	(0.0113)	(0.0097)	(0.0110)
Z2 growth rate*	na†	0.0002*	−0.0002**	−0.0004***	−0.0003**	−0.0003**
		(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
growth rate	−0.035	−3.3455	1.4504*	0.8461	0.7717	−0.6956
		(4.8952)	(0.8011)	(0.8585)	(0.7493)	(0.8098)
Rel. Price of Investment Goods	−0.046**					
Year = 1977?			−0.5088		−0.5661	
			(2.5124)		(1.5158)	
Year = 1982?			8.1445***		6.7653***	
			(2.3096)		(1.7831)	
No. of obs	916	46	153	153	153	153
Chi^2					72.92	18.51
F statistic	22.55	10.28	10.84	7.63		
Prob > Chi^2 (or F)	<0.001	0.0000	0.0000	0.0000	0.0000	0.0024

Notes: Standard errors in parentheses. Dependent variable is savings as a percent of aggregate real income in column 2, and 100 minus personal consumption expenditures as a percent of total personal income in columns 3–7. Estimates for constant term not reported. Growth rate is annual percent change in aggregate GDP. The original Higgins and Williamson model included a control for the relative price of investment goods, which is proxied in columns 5 and 7 using dummy variables for year under the assumption that there would be little cross-sectional variation in this measure in the U.S. data. The Higgins and Williamson model also included a lagged dependent variable to correct for serial correlation, which is not required in the U.S. data. Model in column 3, for 1900, estimated using robust regression and GLS. Z1 and Z2 are calculated using logged population shares, as in the original Higgins and Williamson (1997) model.

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$. †These coefficient estimates for the interaction terms were not provided in Higgins and Williamson (1997).

7. Resolving the Paradox: Are Different Models Needed for DCs and LDCs?

Is there a model that bridges the apparent gap between early and late 20th century U.S. experience (and, by analogy, between age structure effects in LDCs and DCs)? Has behaviour changed, or only appeared to change because of the effects of other factors? And why do various models produce "inconsistent" results for the interaction term between dependency and the growth rate?

If in fact age structure effects vary by level of income, and relative cohort size, the models in the previous section cannot capture those changes since none control for relative cohort size and most exclude income as an independent variable, including it only as the denominator in the dependent variable and thus constraining its coefficient to unity. The one exception appears in Collins (1991)—but Collins does not include an interaction term between dependency and income, nor does she include any control for changing relative cohort size. It is significant, though, that her estimated (positive) coefficient on the dependency rate increases by 50 percent and becomes statistically significant when low-income countries are excluded from her full model (with the savings rate as dependent variable).

A more appropriate model, then, would seem to be one that includes, along with the variables used in previous models, a control for relative cohort size, and some form of income on the right-hand side, together with its interaction with the age structure variables. If the model is estimated in logged form the exact income specification does not really matter, as long as income and total population are included on the right-hand side as separate variables: all forms are mathematically equivalent[12]. Total population is used in the model, although virtually identical results are produced using just the adult population, or the working-age population.

Relative cohort size is specified here using the age groups hypothesized to be least substitutable in the labour force; those aged 20–24 (inexperienced workers) relative to those aged 45–49 (prime-age workers). The results presented in the remainder of this chapter were found to be fairly insensitive to changes in this measure, however.[13] Table 3 presents summary statistics on a range of relative cohort size measures.

Using the Lebergott data, the following basic model (equation (5)) has been estimated:

$$
\begin{aligned}
-\ln \mathrm{PCE}_{S_t} = {} & \varphi_0 + \sum_{j=1}^{J} \varphi_{0j} \ln(p_j) + \varphi_1 \ln \mathrm{RCS} + \varphi_2 \ln y_{St} + \ln y \sum_{j=1}^{J} \varphi_{2_j} \ln(p_j) \\
& + \varphi_3 \ln P_{S_t} + \varphi_4 g_{S_t} + v_S + u_{S_t}
\end{aligned} \tag{5}
$$

[12] Although the form of the income variable that is interacted with the age structure variables will change the *estimated coefficient on the interaction term*, it will have no effect on the estimated *total* effect of each age share on the dependent variable, at various levels of income, when the model is estimated in logged form.

[13] That is, although the level of statistical significance varies among the different measures, the estimates of overall effect of changing age structure, as presented in Figures 4–6, are virtually unaffected by changes in the ratio used. These additional results are available from the author on request.

Table 3. Summary statistics on a range of relative cohort size variables, at the state and national levels. The ratio of 20–24/45–49 year olds was used in this analysis.

Variable (by Year)	**Obs**	**Mean**	**Std. Dev.**	**Min**	**Max**	**U. S. actual levels**
			Ratio 15–24/total population			
1900	49	0.20	0.02	0.17	0.23	0.20
1929	49	0.19	0.02	0.16	0.22	0.18
1970	51	0.19	0.01	0.16	0.21	0.18
1977	51	0.20	0.01	0.17	0.23	0.19
1982	51	0.18	0.01	0.16	0.19	0.18
U.S. averages:						
1900–1925	26	0.19	0.01	0.18	0.20	
1926–1950	25	0.18	0.01	0.15	0.18	
1951–1975	25	0.16	0.02	0.13	0.19	
1976–1999	24	0.16	0.02	0.14	0.19	
			Ratio 15–24/25–59			
1900	49	0.49	0.10	0.32	0.71	0.48
1929	49	0.43	0.08	0.30	0.63	0.41
1970	51	0.47	0.04	0.37	0.58	0.46
1977	51	0.47	0.04	0.39	0.59	0.45
1982	51	0.40	0.03	0.34	0.47	0.40
U.S. averages:						
1900–1925	26	0.44	0.03	0.40	0.48	
1926–1950	25	0.38	0.03	0.31	0.41	
1951–1975	25	0.38	0.07	0.29	0.46	
1976–1999	24	0.35	0.06	0.28	0.45	
			Ratio 15–24/45–54			
1900	49	2.27	0.46	1.50	3.32	2.25
1929	49	1.75	0.37	1.11	2.63	1.70
1970	51	1.67	0.20	1.27	2.30	1.60
1977	51	1.94	0.25	1.47	2.72	1.82
1982	51	1.88	0.14	1.57	2.35	1.83
U.S. averages:						
1900–1925	26	1.98	0.19	1.73	2.25	
1926–1950	25	1.53	0.15	1.24	1.75	
1951–1975	25	1.38	0.21	1.13	1.75	
1976–1999	24	1.50	0.32	1.02	1.87	
			Ratio 20–24/45–49			
1900	49	2.05	0.38	1.53	2.98	2.06
1929	49	1.57	0.30	1.04	2.22	1.52
1970	51	1.52	0.21	1.11	2.35	1.49
1977	51	1.92	0.25	1.46	2.74	1.85
1982	51	1.99	0.18	1.60	2.45	1.95

Table 3. (*Continued)*

Variable (by Year)	Obs	Mean	Std. Dev.	Min	Max	U. S. actual levels
U.S. averages:						
1900–1925	26	1.82	0.21	1.52	2.06	
1926–1950	25	1.41	0.09	1.23	1.53	
1951–1975	25	1.24	0.23	1.00	1.72	
1976–1999	24	1.49	0.40	0.93	1.99	
		Ratio 20–22/45–49				
1900	49	1.23	0.24	0.90	1.82	1.25
1929	49	0.95	0.19	0.62	1.37	0.92
1970	51	0.92	0.13	0.65	1.38	0.90
1977	51	1.17	0.15	0.89	1.70	1.14
1982	51	1.21	0.11	0.95	1.46	1.18
U.S. averages:						
1900–1925	26	1.10	0.12	0.92	1.25	
1926–1950	25	0.85	0.06	0.72	0.93	
1951–1975	25	0.77	0.15	0.60	1.06	
1976–1999	24	0.90	0.24	0.58	1.21	

where S is state (from 1 to 51); t is year (1900, 1929, 1970, 1977, and 1982); PCE is personal consumption expenditure as a share of total income, with its inverse treated as savings; RCS is relative cohort size, the ratio of population aged 20–24 relative to those aged 45–49; p_j is the age group j share of total population; y is real per capita personal income; P is total population; and g is the growth rate of real GDP.[14] In addition, four alternative formulations were tested, including controls for:

1. the percent of a state's population that is foreign-born, and the percent living in urban areas, both of which can be expected to increase consumption's share of income [equation (6)];
2. interactions between the growth rate and age structure variables, as in the Fry and Mason formulation [equation (7)];
3. dummy variables for year, to capture any period effects [equation (8)]; and
4. a model containing all of the above variables [equation (9)].

---

[14] Total population is also needed in the model to control for the fact that the denominator is lost in calculating Zs with logged population shares. A reformulation of equation (3), substituting s_j/P for p_j, where s_j is the absolute number in age group j and P is total population, indicates that if $\ln(s_j/P)$ is used in this equation in place of s_j/P, the total population terms cancel, leaving a variation on Z_n based on absolute age group sizes, rather than age group shares:

$$X_{i,t} = \gamma_{0,i} + \gamma_1 X_{i,t-1} + \gamma_2 g_{i,t} + \gamma_3 RPI_{i,t} + \gamma_4 Z_{1_{i,t}} + \gamma_5 Z_{2_{i,t}} + \gamma_6 g_{i,t} Z_{1_{i,t}} + \gamma_7 g_{i,t} Z_{2_{i,t}} + u_{i,t}$$

Apparently the Higgins (1998) and Higgins and Williamson (1997) formulations both used logged population shares. Higgins (p. 348, footnote 7) notes that "the (coefficients on the age shares) must be restricted to sum to zero because the population age shares sum to unity, and thus are collinear with the intercept term", suggesting that he is working with unlogged population shares. However, in all of his table notes (as, for example, on p. 357) he explains that "the age distribution coefficients represent the change in the dependent variable associated with a unit change in the corresponding log population age shares."

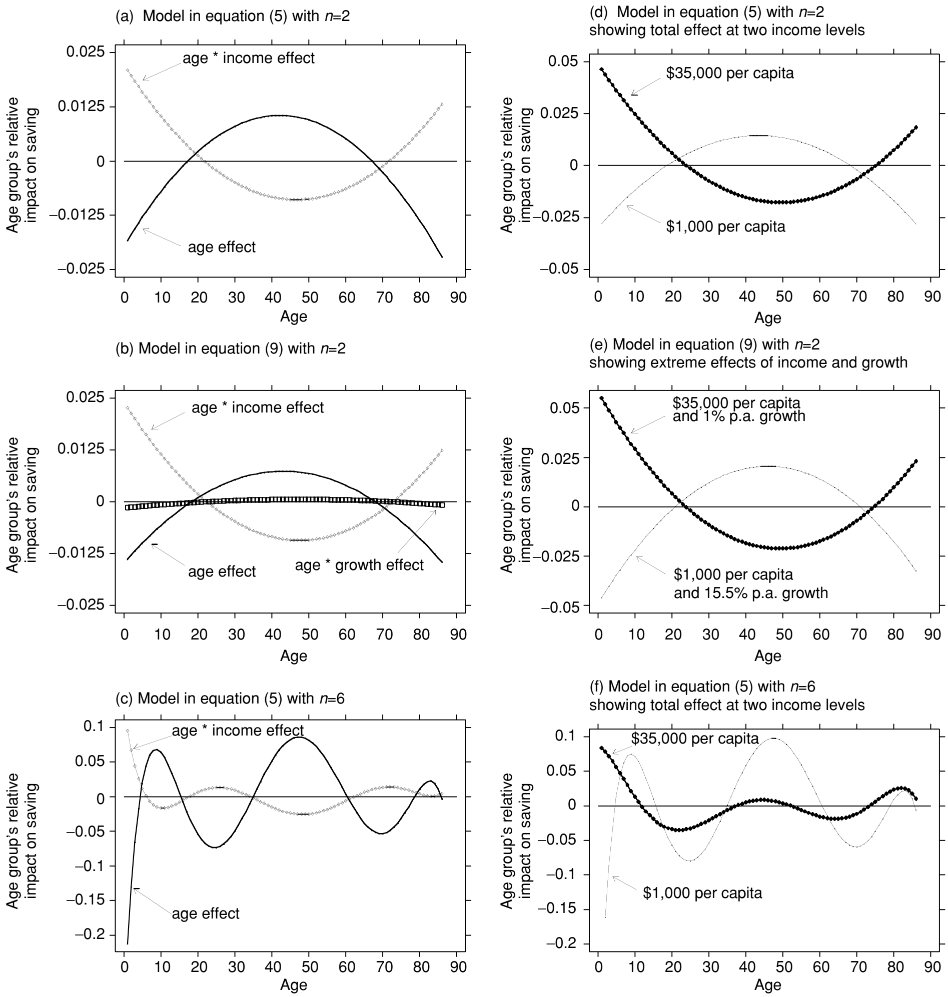

Figure 4. Estimating age-share coefficients, using models in columns 2 and 6 of Table 4.
Left: Panels 4a–c presents estimated effects of each age group on the intercept (coefficients on the age shares themselves) and slope (coefficients on the interaction terms between age shares and income and growth) of an equation estimating savings as a share of total personal income using Lebergott (1997) state-level U.S. data for 1900, 1929, 1970, 1977, and 1982.
Right: Panels 4d–f present estimates of the net effect of age structure at high and low levels of per capita income ($1,000 and $35,000, in year 1999 dollars, using the minimum and maximum values in the data). Panel 4e also incorporates the esitimated effects of growth on the age structure coefficients, and thus might be thought to present "extreme" cases: a low-income country ($1,000 per capita) experiencing rapid growth (15.5 percent based on the maximum in the data) and a high-income country ($35,000 per capita) experiencing low growth (1 percent p. a.). The coefficients in these three graphs represent elasticities. For example, Figure 4e suggests that a 10 percent

Although each of these additional controls produces statistically significant coefficients, and increases the overall explanatory power of the model, none have any substantial effect on the estimated level and pattern of age structure effects over time, as illustrated in Figures 4–6, and explained in the next section of this chapter.[15]

But what should be the degree of the polynomial used to represent age structure effects? Higgins and Williamson (1997) and Higgins (1998) assumed quadratic and cubic forms, no doubt with the life-cycle model in mind. Similarly Fair and Dominguez (1991) used a second-degree polynomial, expressly in the expectation of a life-cycle pattern of effects. But as emphasized earlier there is no theoretical reason to expect the life-cycle pattern to dominate at the aggregate level: once interhousehold effects are included we are no longer estimating a behavioural model.

The testing procedure adopted here, with the models indicated above, is that recommended in Judge (1985), i.e., starting from a higher degree polynomial than might be thought appropriate ($n = 8$) and sequentially adding restrictions. Although this procedure has been shown to be problematic in time series data when *both* the degree of the polynomial (n) and number of lags must be estimated (Terasvita 1976; Schmidt and Mann 1977; and Schmidt and Waud 1973), it is reliable when the number of lags (in this case, the maximum age) is known (Godfrey and Poskitt 1975; Harper 1977; Frost 1975; Schmidt and Sickles 1975). The results of these tests are presented in Table A.1, where it can be seen that only the quadratic form ($n = 2$) is unambiguously indicated. Only at that level are the t and Wald chi-squared statistics significant in all tests performed. The t-statistics on the Z's and their interactions with income are significant for $n = 6$ in equation (5) and $n = 5$ in equation (9), as well—but Wald tests of the highest coefficients in the fifth- and sixth-degree polynomials are singular, suggesting that the model is over fitted when $n > 2$.[16] However, because the overall explanatory power of the sixth-degree polynomial (as indicated by R^2 and Mean Squared Error (*MSE*)) is in general greater than that of the quadratic, the results presented here will first examine the quadratic model, and then the effect of adopting the sixth-degree model, on the pattern of estimated age share coefficients and their ultimate effect in simulations at the national level.

Figure 4. (*continued*)
increase in the population share of 1-year-olds would generate a 0.5 percent decline in the savings rate in a low-income high-growth economy, whereas the same increase would generate a 0.6 percent increase in the savings rate in a high-income low-growth economy.
The top four panels illustrate the patterns of age share coefficients obtained using a quadratic polynomial ($n = 2$), while the bottom two illustrate the patterns obtained using a sixth-degree polynomial.

[15] Similarly, even changes in the growth rate and total population size have virtually no effect on the estimated pattern of age structure effects, although excluding the growth rate altogether reduces the estimated level of age structure effects by about one-third.

[16] Results from earlier estimates of a model without any control for relative cohort size are presented in Macunovich (2001). Without that additional control, fourth- (and sixth-) degree polynomials are required to fit the data.

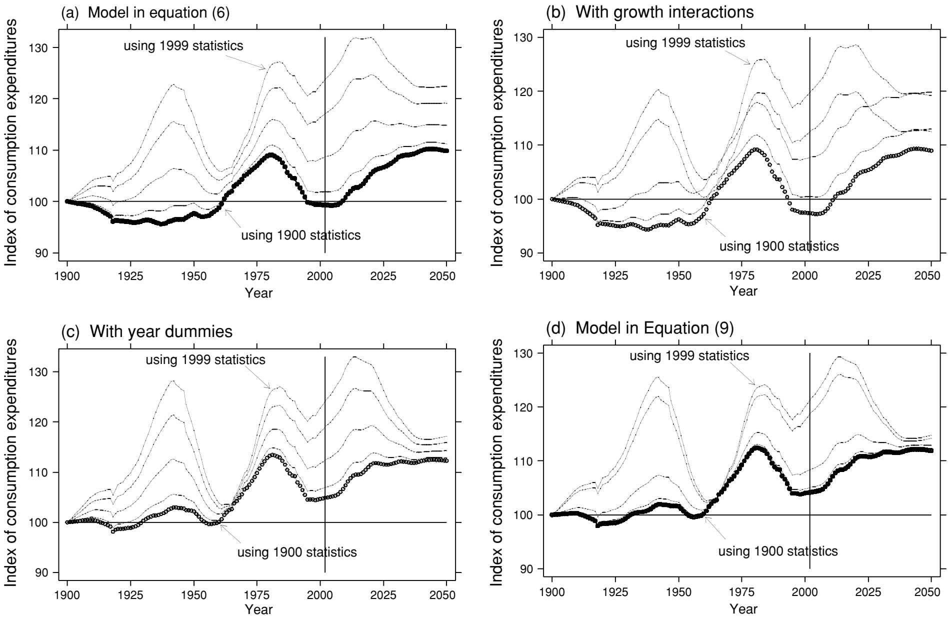

Figure 5. Simulated effect of changing age structure on absolute personal consumption expenditures in the United States, holding everything except age structure constant. Each curve holds real income, total population size, and GDP growth constant at the levels experienced in either 1900, 1925, 1950, 1975, or 1999. The four models, used to prepare the simulations are presented in columns 3–6, respectively, in Table 4. There are strong similarities among the results from all models, especially in the marked declines from 1930 to 1960, increase from 1960 to 1980, and declines again between 1980 and the mid 1990. Rates of change in these indexes are presented in Figure 6.

Including income and its interaction with the population shares in this way, together with a relative cohort size measure, creates a model that explains the apparent reversal in age structure effects between low- and high-income economies produced in earlier models, as demonstrated in Tables 4 and 5, and in Figure 4. Column 2 in Table 4 presents the estimates using $n = 2$ in the basic model set out in equation (5), and columns 3–6 illustrate the effect of successively adding controls for percents urban- and foreign-born, interaction terms between age structure and the growth rate, and year dummies. Table 5 presents results of model estimates designed to illustrate the significance of the age structure variables in alternative formulations, again in the quadratic model; the coefficients on the age structure variables, their interaction with income, and on relative cohort size remain highly significant in all formulations.

The top four panels in Figure 4 illustrate the patterns of age structure coefficients estimated in models (5) and (9), with $n = 2$, i.e., with age share coefficients constrained to lie along

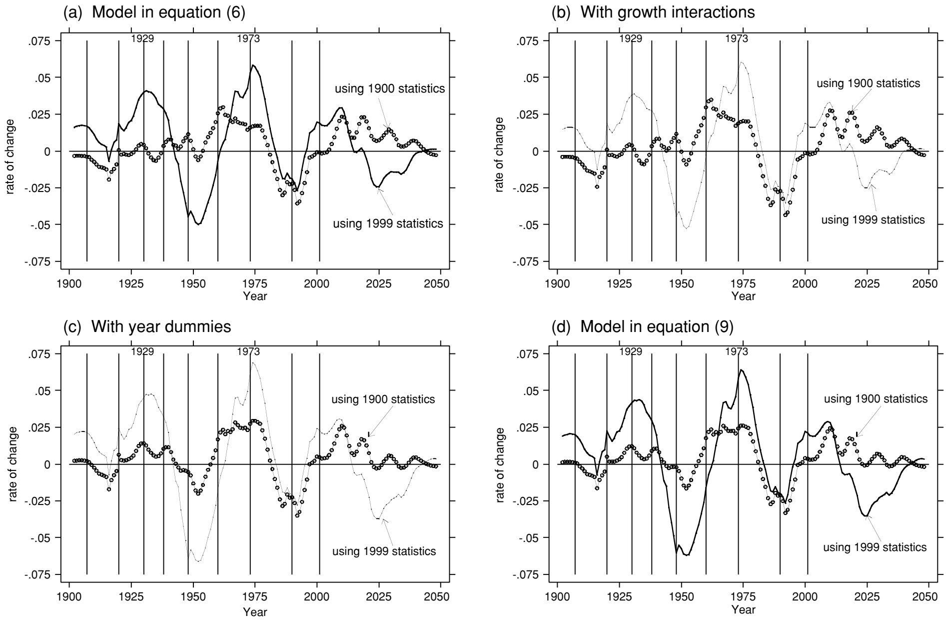

Figure 6. Rates of change in indexes of simulated absolute personal consumption expenditures generated by age structure changes in the United States as presented in Figure 5. The four models used to prepare the simulations are presented in columns 3–6, respectively, in Table 4. There are strong similarities among the results from all models, in terms of the timing of peaks and troughs. The vertical lines indicate peaks in consumption expenditure growth that coincide with business cycle peaks. These account for 9 of the 21 20th century business cycle peaks identified by the NBER, including some of the most significant, such as those in 1929 and 1973.

a quadratic polynomial. Figures 4a and 4b illustrate the patterns of age share coefficients estimated in the two models presented in columns 2 and 6 (equations (5) and (9)), respectively, of Table 4. Each graph illustrates two separate sets of age share coefficients—the effect on the intercept (coefficients on the age shares themselves) and slope (coefficients on the interactions between age shares and per capita income) of a savings equation. The simple age share effects conform to the standard "hump-shaped" life-cycle pattern of saving and dis-saving, but the hump is inverted for the coefficients on the interaction terms with income, suggesting a shift in the total effect of age structure as real per capita income rises.

It is difficult to conceptualize the total effect of age structure from the panels on the left of Figure 4, however, since the total effect depends on the magnitude of the interaction effects, which in turn depends on level of real per capita income. Thus, the panels on the right in Figure 4 illustrate the combined effects of age structure in a savings equation, at two different levels of per capita income—$1,000 and $35,000 in year 1999 dollars. Figure 4e

Table 4. Full regression results for the savings rate as a function of age structure variables, using Lebergott state-level data for the U.S. in 1900, 1929, 1970, 1977, and 1982. Columns 3–6 in this table correspond to the four models used in Figures 5 and 6. Columns 2 and 6 in this table were used as the basis for simulations reported in Figure 4.

	Equation (5)	Equation (6)	w/growth interactions	w/year dummies	Equation (9)
Relative cohort size	−0.125***	−0.134***	−0.161***	−0.140**	−0.132**
	(0.043)	(0.040)	(0.041)	(0.063)	(0.063)
Z1	0.00143***	0.00193***	0.00172***	0.00116***	0.00103***
	(0.0005)	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Z2	−1.70e-05***	−2.11e-05***	−1.88e-05***	−1.33e-05***	−1.20e-05***
	(4.90e-06)	(4.60e-06)	(4.70e-06)	(4.30e-06)	(4.30e-06)
Z1 * income/capita	−0.00133***	−0.00141***	−0.00141***	−0.00141***	−0.00140***
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Z2 * income/capita	1.43e-05***	1.45e-05***	1.46e-05***	1.47e-05***	1.47e-05***
	(2.50e-06)	(2.30e-06)	(2.40e-06)	(2.20e-06)	(2.20e-06)
Z1 * growth rate			9.38e-05**		8.75e-05**
			(4.59e-05)		(4.11e-05)
Z2 * growth rate			−1.10e-06**		−9.00e-07**
			(5.00e-07)		(4.00e-07)
real income/capita	0.272***	0.226***	0.227***	0.428***	0.428***
	(0.05)	(0.048)	(0.048)	(0.065)	(0.067)
growth rate of GDP	−0.015***	−0.014***	−0.026***	−0.008***	−0.009
	(0.003)	(0.003)	(0.008)	(0.003)	(0.007)
total population	0.0095	0.025***	0.027***	0.026***	0.027***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
% urban		−0.118***	−0.115***	−0.116***	−0.126***
		(0.023)	(0.024)	(0.022)	(0.023)
% foreign born		−0.018***	−0.019***	−0.038***	−0.034***
		(0.007)	(0.007)	(0.009)	(0.009)
Year = 1929?				0.010	0.020
				(0.023)	(0.024)
Year = 1970?				−0.252***	−0.232***
				(0.060)	(0.063)
Year = 1977?				−0.276***	−0.263***
				(0.065)	(0.066)
Year = 1982?				−0.209***	−0.197***
				(0.072)	(0.073)
Constant	−0.564***	−0.698***	−0.666***	−1.025***	−1.025***
	(0.127)	(0.120)	(0.122)	(0.133)	(0.138)
sigma_u	0.04029	0.035102	0.03435	0.0384	0.037827
sigma_e	0.075900	0.074200	0.073396	0.062970	0.062216
rho	0.219913	0.182874	0.179700	0.271119	0.269896
R^2: within	0.7599	0.7702	0.7782	0.8319	0.8363
between	0.1843	0.4130	0.4324	0.4622	0.4631
overall	0.6320	0.6996	0.7082	0.7569	0.7597
Wald chi^2	546.11	650.90	673.23	941.78	955.92
Prob > chi^2	0.0000	0.0000	0.0000	0.0000	0.0000

Notes: Dependent variable is − ln (personal consumption expenditures/total personal income). 248 observations. All variables except growth rate used in logged form. Standard errors in parentheses. Estimated as random effects models using GLS to control for heteroscedasticity.

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$.

Table 5. Partial models of the savings rate as a function of age structure variables, using the Legbergott state-level U.S. data. Here it can be seen that the interaction terms between age structure and the growth rate are not significant (column 6) until a control for relative cohort size is introduced (column 7), and even then these interactions are significant only at the 10 percent level. Columns 4 and 6 in Table 4 indicate that these variables become much more significant when income interactions are added to the model. The relative cohort size variable remains highly significant (in results not presented here) even when no other variables are included in the model.

	Equation (5)					
Relative cohort size		−0.129*** (0.033)		−0.125*** (0.043)		−0.235*** (0.040)
Z1			0.00205*** (0.0004)	0.00143*** (0.0005)	−0.00041* (0.0002)	−0.00115*** (0.0002)
Z2			−2.26e-05*** (4.64e-06)	−1.70e-05*** (4.91e-06)	3.66e-06 (2.28e-06)	1.09e-05*** (2.45e-06)
*Z1 * income/capita*			−0.00154*** (0.0002)	−0.00133*** (0.0002)		
*Z2 * income/capita*			1.61e-05*** (2.52e-06)	0.0000143*** (2.55e-06)		
*Z1 * growth rate*					8.02e-05 (5.39e-05)	9.17e-05* (5.08e-05)
*Z2 * growth rate*					−8.28e-07 (5.67e-07)	−1.04e-06* (5.35e-07)
real income/capita	0.170*** (0.010)	0.168*** (0.010)	0.199*** (0.044)	0.272*** (0.050)	0.205*** (0.021)	0.244*** (0.020)
growth rate of GDP	−0.007*** (0.003)	−0.007*** (0.003)	−0.012*** (0.003)	−0.015*** (0.003)	−0.008 (0.008)	−0.028*** (0.008)
total population	−0.004 (0.008)	−0.006 (0.008)	0.010 (0.008)	0.010 (0.007)	0.003 (0.008)	0.009 (0.007)
constant	−0.156*** (0.057)	−0.064 (0.061)	−0.444*** (0.122)	−0.564*** (0.127)	−0.372*** (0.103)	−0.413*** (0.096)
sigma_u	0.051234	0.053423	0.043497	0.04029	0.040087	0.034738
sigma_e	0.087571	0.083229	0.078360	0.075900	0.086738	0.080427
rho	0.255003	0.291787	0.235552	0.219913	0.176003	0.157225
R^2: within	0.6700	0.7045	0.7457	0.7599	0.6716	0.7216
between	0.0583	0.0519	0.1766	0.1843	0.1451	0.2400
overall	0.5323	0.5388	0.6238	0.6320	0.5555	0.6031
Wald chi^2	371.24	418.65	527.44	546.11	363.67	441.95
Prob > chi^2	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000

Notes. Dependent variable is − ln (personal consumption expenditures/total personal income). 248 observations. All variables except growth rate used in logged form. Standard errors in parentheses. Estimated as random effects models using GLS to control for heteroscedasticity.

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.10$.

compares a low-income ($1,000 per capita) but high-growth (15.5 percent p.a.) economy, and a high-income ($35,000 per capita) but low-growth (1 percent p.a.) economy, consistent with the standard model of economic convergence. In Figures 4d and 4e the pattern of age structure effects at low incomes, as in LDCs, is “hump” shaped as in the life-cycle model, but at levels of income typical of those in industrialized nations *the pattern is inverted*. This is the same effect obtained in Table 2 with the Higgins and Williamson (1997) model—but there it was necessary to estimate two separate models to fit the two sets of data, for low- and high-income periods.

The coefficients presented in Figure 4 represent elasticities. For example, Figure 4d suggests that a 10 percent increase in the population share of 1-year-olds would produce a 0.25 percent drop in the savings rate at the aggregate level in LDCs, but a 0.5 percent *increase* in the savings rate in high-income economies. Figure 4e indicates that although the *overall pattern* of age share coefficients is not changed in moving from equation (5) to equation (9), the *magnitude* of the total effect increases. In Figure 4e a 10 percent increase in the population share of 1-year-olds produces a 0.5 percent drop in the savings rate at the aggregate level in the low-income high-growth economy, but a 0.6 percent *increase* in the savings rate in the high-income low-growth economy.

For those who find a complete inversion of the life-cycle savings pattern at higher levels of income difficult to accept, it is perhaps helpful to consider how the pattern of coefficients changes as the degree of the polynomial increases. As mentioned earlier, there is some support in the statistical tests for the sixth-degree polynomial (in terms of higher R^2 and significant t-statistics on the highest Z’s, although Wald tests are singular). What is the pattern of age share coefficients produced using a sixth-degree polynomial? This is demonstrated in the bottom two panels of Figure 4, where (in panel 4f) the standard hump-shaped life-cycle pattern is retained for ages 15–75 even at the higher income level, although with a considerably reduced amplitude.

Although the pattern of effects estimated here is more complex than the standard life-cycle model, it is not inconsistent: it simply adds a new wrinkle, with respect to the effect of children once we have controlled for the ages and spending patterns of their parents. In addition, these parameters do not necessarily reflect individual behaviour as it might be observed at the household level; rather they estimate the behaviour induced by various age groups at the societal level, including all interhousehold and intergenerational spending. Given the complexity of the effects estimated here, it is hardly surprising that studies over the years have produced so many different estimates of the effect of dependency on savings and growth.

8. Might These Effects Be Significant in Terms of Macroeconomic Fluctuations?

Returning to the question at the opening of this chapter, what impact has changing age structure had on actual patterns of consumption in the United States throughout the 20th century, as estimated by this model? The model results presented in the previous section are all highly significant *statistically*, but how substantively significant are they in *real world* terms? Whereas for the reasons discussed earlier it would be inappropriate to attempt to answer that question at the national level using parameters obtained in a micro level analysis, this analysis has provided macro level parameters, measured at the state level that *can* be applied at the national level. The goal is to use them to simulate the pattern of PCE over time in the United States at the national level that would have resulted if nothing other than age structure had changed: i.e., holding constant the total population size, income level, and GDP growth rate.

But holding income, population size and growth constant makes it very difficult to illustrate the full effect of changing age structure, both directly and as it operates through income and the growth rate. In order to estimate the full effect of changing age structure, each of the graphs in Figure 5 (which are based on the models in columns 3–6 of Table 4, moving from equations (5) to (9) with $n = 2$) presents the results of five different simulations for the United States. In each one, all variables except age structure are held constant, but at a different level—one, at the levels observed in 1900, another at 1925 levels, and three others at 1950, 1975, and 1999 levels, respectively. They trace out undulating "ribbons" of effects over the century on total PCE, with a maximum range of about 25 percent using the 1999-level characteristics, but applied to age structure changes that actually occurred in the 1960 to 1980 period.

It is worth emphasizing the meaning of this figure. If, in 1999, the age structure had been that observed in 1980, when the largest baby boom cohorts were about 20 years old—but with all other factors including income and population size held constant—the level of PCE would have been *25 percent higher* than the expenditures that would have occurred with a 1960 U.S. age structure, when birth rates were at their peak and the share of 20-year-olds was very low. This swing is virtually unaffected by the different formulations of the model that are presented in columns 3–6 of Table 4.

Despite the difference in levels among the five curves in each graph in Figure 5, they all have in common extremely strong growth between 1960 and 1980, and then decline after 1980, as a result of the baby boom's passing from childhood to young adulthood. It is outside the scope of this study to say whether this is a "good" or "bad" effect; it implies strong growth in aggregate demand, but could obviously result in declining savings rates unless per capita incomes rise correspondingly. And, of course, they did rise dramatically in the 1960s and early 1970s—assisted at least in part by the rising domestic aggregate demand. As Higgins and Williamson (1997) and Higgins (1998) emphasize, in open economies an increase in the share of domestic income devoted to consumption need not result in a decline in overall investment and hence a decline in productivity and real growth. These researchers trace a pattern of increasing age-structure-induced investment supported by an increasingly negative current account balance, culminating in rising productivity and growth in real per capita output (and a reversal to a positive current account balance) as children become productive young adults.

The similarity in the estimated effects in the four panels of Figure 5 is emphasized if we examine the *rates of change* in consumption expenditures, as in Figure 6, rather than the levels shown in Figure 5. Each of the four panels in Figure 6 illustrate rates of change in consumption expenditures using the 1900 and 1999 values for income, total population size, and growth rate. The similarities in the four panels are striking, with peaks occurring

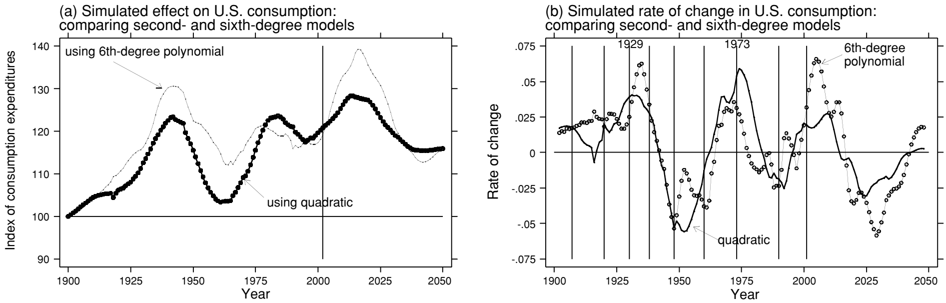

Figure 7. The effect of moving from a quadratic to a sixth-degree polynomial, on simulated impact of changing age structure on U.S. personal consumption expenditures.

in virtually the same years regardless of model formulation or levels of nonage-structure variables. The vertical lines in Figure 6 mark peaks in the growth rate of age-structure-induced consumption expenditures that coincide with business cycle peaks during the 20th century. This coincidence of age-structure-induced peaks and business cycle peaks occurs for at least 9 of the 21 business cycle peaks identified in the 20th century by the NBER, including several of the most significant, such as those in 1929 and 1973.

And finally, there is the question of sensitivity to model formulation. Although statistical tests indicate that anything higher than a quadratic polynomial for the age structure coefficients probably overfits the data, it is nevertheless worthwhile to examine the effect of moving from a quadratic to a sixth-degree polynomial in estimating age structure effects at the aggregate level. Figure 7 presents a comparison, using 1999 levels of income, total population size, and growth rate, of simulated effects using the age share coefficients estimated with two different polynomials in equation (5): $n = 2$ and $n = 6$. The two patterns are remarkably similar, in terms of the overall sinusoidal pattern, although differences between the two appear to be greatest in the most recent period when age structure changes have been most dramatic, with considerably wider swings over time estimated using the sixth-degree polynomial.

9. Conclusions

As expected, the results of this analysis support the hypothesis that the dependency effect changes with income level; children in lower income economies have an overall positive (negative) effect on the consumption (savings) rate, but as per capita income rises that effect is reversed. Children in high-income areas appear to motivate increased saving—or at least relatively lower levels of consumption—on the part of other members of the population, whether through precautionary motives or to finance bequests. Perhaps when infant and child mortality rates are low, and we turn from a desire for "quantity" to "quality" in our offspring, children make us more forward-looking: they provide a reason to place increased weight on future consumption. If so, and if indeed, the savings rate has declined in the United States over the past two decades, perhaps that decline occurred because there are relatively fewer children to save *for*, which would be ironic given the negative emphasis often placed on children in the dependency literature.

But despite this shifting effect, the estimates presented here suggest that changes in age structure—especially those due to the baby boom—have had a major impact on the U.S. economy, probably contributing a huge boost in the 1960–1980 period, but then a considerably weakened effect after 1980. Boomers and their children are likely to exert another strong positive effect in the first part of the 21st century, based on the parameters in the model and current Census Bureau population projections.

In addition, exploratory work on nations around the world from 1950 to 2000 indicate a strong correlation between changing age structure and the incidence of "financial crises" over the past 20 years in developing nations, as a type of delayed fallout from the demographic transition.[17] Perhaps the boom in consumption expenditures generated by increasing young adult cohorts contributes to "irrational exuberance" on the part of investors and lenders—which then collapses when growth in the size of those cohorts switches to decline.[18]

But obviously this analysis is only a beginning, given this "new" technique for estimating age structure effects that was introduced by Fair and Dominguez (1991). It suggests a very complex pattern of age structure effects that is consistent with the life-cycle model, but at the same time generates a much wider range of dependency effects than has been contemplated up to now in the literature. It is hoped that the results here will generate additional interest in using full age distributions to examine aggregate age structure effects, and allowing for the effects of interhousehold transfers on aggregate patterns of age-specific effects.

Appendix: Data Sources

The data on consumption expenditures which were used in the analysis were prepared by Stanley Lebergott (1997), in a painstaking and well-documented effort which produced PCE at the state level, broken down into the 100+ sub-categories itemized in the national income accounts, for the years 1900, 1929, 1970, 1977, and 1982, as well as annual data at the national level for the years 1900–1993. The same source provides personal income figures at the state and national levels, for the same dates.

As Lebergott describes, his data were developed chiefly from

> the Censuses of Retail Trade, Services, Housing, Government and Population. Each Census drew on an enormous sample of knowledgeable respondents. For the most part the reports rest on detailed and original records rather than fleeting and vagrant consumer memories. Thus the 1977 Census of Retail Trade collected data from firms with over 85% of all retail sales. The

[17] A discussion of this potential relationship is presented in Macunovich (2002).

[18] This pattern is supported by findings in Higgins and Williamson (1997) and Higgins (1998), who find that GDP growth and investment spending in Pacific Rim nations was substantially increased as postdemographic transition population "bulges" entered productive adult life.

> 1977 Census of Services relied on direct reports for about 70% of all services in scope. The 1980 Census of Housing collected rent and value data from over 95% of the population. (p. 71)

His state estimates

> uniformly began with estimates, by specific item, for the United States as a whole . . . Each U.S. total was distributed among the states by an allocator, and usually checked against another allocator. Had estimates been directly made for each state, they would not necessarily add to an adequate U.S. total. More important, comparisons against the per capita average for the U.S. as well as nearby states, permit some judgment as to whether a state estimate falls outside reasonable limits.

It is important to note that in no case were the allocators used by Lebergott based on age distributions within the population. Rather, they were derived from various census data on production and expenditures, as well as distributions of workers by occupation and service income. The following passages, taken from his documentation for 1900, illustrate the methodologies employed by Lebergott in allocating expenditures at the state level.

> For such major items as food, clothing, furniture, and lighting we utilize the 1901 expenditure survey of 25,440 families by the U.S. Commissioner of Labour. Our individual state averages for these individual items were checked by regressing them against relevant occupation counts times average nonagricultural service income. (p. 92) For food off-premise the result was then checked against the Population Census count of persons engaged in food retailing: merchants and dealers (excl. wholesale) in groceries and produce, hucksters and peddlers, butchers, bakers, and confectioners. (p. 93) A U.S. total of meals and beverages "was allocated by the number of persons in specified occupations (hotel keepers, bartenders, restaurant keepers, saloon keepers, and waiters) times the average service income per worker in the state." (p. 93) Food furnished to employees was allocated using two series: "One was the aggregate expenditures on farm labour reported by farmers. The other was (a) average monthly wages without board, minus average wages with board, divided by (b) farm wages without board." (p. 94)
>
> The value of dairy products consumed on farms was used to allocate the U.S. total for food produced and consumed on farms by farm operators. (p. 94)
>
> Clothing expenditures per capita in thirty-three states can be derived from the survey by the Commissioner of Labour. The intra-regional variation shown by these figures seemed unreasonably great, and was probably a reflection of sampling variability. We therefore averaged the per capita figures within each of eight regions. These averages were then tested by correlating them with per capita expenditures given by multiplying the occupation count for two groups of merchants and dealers (clothing and men's furnishings plus dry goods, fancy goods, and notions) times the service income per worker . . . (p. 95)

POPULATION DATA

The population data were taken from hard copy and electronic files provided by the Bureau of the Census. These data were on occasion available only by 5-year age group: in such cases, single years of age were estimated as one-fifth of each corresponding 5-year age group (although econometric results were found to be unchanged using simple 5-year age groups).

1900

Data at the state level on total population and population by 5-year age group through age 34, and 10-year age group 35–64, were taken from the 1900 Census of Population, Volume 1, Part 1, Table 3, pp. 110–111.

These data were supplemented by percentage distributions of state populations by 5-year age group through age 84, provided for 1900 in the 1930 Census of Population, Volume II: General Report, Statistics by Subject, Table 25, pp. 660–668.

1929

Data for the year 1930, as provided in the 1930 Census of Population, were used for 1929. Data at the state level on total population and population by 5-year age group through age 84 were taken from the 1930 Census of Population, Volume II, General Report, Statistics by Subject, Table 24, pp. 610–658.

1970

Data at the state level on population aged 0–2, 3–4, 5–13, by single year of age from 14–24, and for age groups 25–29, 30–34, 35–44, 45–54, 55–59, and 65+ were available in electronic form for the year 1970 in file e7080sta.txt from the Bureau of the Census website *http://www.census.gov/population/estimates/state*

The 1970 estimates in this file are consistent with those published in Current Population Reports Series P-25, No. 998. These electronic data were supplemented using total population and population by 5-year age group through age 84 taken from the 1970 Census of Population, General Population Characteristics, Volume I, Section I, Part 1, Table 62, pp. 297–309.

1977

Two alternative methods were used to prepare population by single year of age for 1977:

1. Data for 1980 are available at the state level by single year of age to 84 (based on the 1980 Census), and there are significant discrepancies between the Census Bureau's population estimates for the late 1970s, and the 1980 census. These discrepancies, taken together with the fact that the 1977 data are too aggregated in certain age ranges (see point 2 following), led to an effort to "backdate" the 1980 data by 3 years, to be used in place of the questionable 1977 figures. These backdated 1980 population figures were assumed to be preferable for calculating population age shares.
2. Data at the state level on population aged 0–2, 3–4, 5–13, by single year of age from 14 to 24, and for age groups 25–29, 30–34, 35–44, 45–54, 55–59, and 65+ were available in electronic form for the year 1977 in file e7080sta.txt from the Bureau of the Census website
   http://www.census.gov/population/estimates/state
   The 1977 estimates in this file are consistent with those published in Current Population Reports Series P-25, No. 998. Because there is no additional source of data for 1977, as there is for 1970, to break down the larger age aggregates (35–39 from 35–44, 45–49 from 45–54, 5–9 from 5–13, and 5-year groups above 64), national patterns

Table A.1. Coverage in each decade.

	1900–1929	1930–39	1940–49	1950–59	1960–79
Age	**0–75+**	**0–75+**	**0–85+**	**0–85+**	**0–85+**
Resident population	Yes	Yes	Yes	Yes	Yes
AK and HI?	excluded	excluded	excluded	included	included
AF overseas?	excluded	excluded	included	included	included
Reference CPR P-25	#311	#11	#311	#311	#519 and #917

for the year 1977 were used within these age groups. These estimated 1977 figures were used as a check on the backdated 1980 figures: it was found that there is a close correspondence.

1982
Data at the state level on total population and population by single year of age through age 84 on July 1, 1982 were taken from computer files downloaded from the Bureau of the Census website
http://www.census.gov/population/www/estimates/st_stiag.html

1900–1979 US data
These data were taken from Lotus files of population by single year of age, provided on diskette by the Bureau of the Census (Kevin Deardorff, Population Division, at 301-763-7950).

1980–1989 US data
(Resident population plus Armed Forces overseas, including AK and HI, age 0–100+ and total, as of July in each year): Files e8081pqi.txt through e8990pqi.txt from Bureau of the Census website
http://www.census.gov/population/www/estimates/nat_80s_detail.html

1990–1998 US data
(Resident population plus Armed Forces overseas, including AK and HI, age 0–100+ and total, as of July in each year): Files e9090pmp.txt through e9898pmp.txt from Bureau of the Census website
http://www.census.gov/population/www/estimates/nat_90s_2.html

Method Used to Estimate Coefficients on Population Age Shares

In an unconstrained model, population age shares—if assumed to affect the intercept of the consumption equation—would enter the equation for personal consumption expenditures as

$$\sum_{j=1}^{J} \varphi_j p_j \tag{A.1}$$

where p_j is the share of total population represented by age group j; and φ_j is the coefficient to be estimated. Since the population data used here are available in single

Table A.2. Statistics used to determine appropriate degree of polynomial (*n*).

Degree of polynomial:	**8**	**7**	**6**	**5**	**4**	**3**	**2**
A. Equation (5)							
R-sq: within	0.8142	0.8147	0.8170	0.8034	0.7972	0.7867	0.7599
between	0.2404	0.2409	0.2363	0.1858	0.1750	0.1802	0.1843
overall	0.6913	0.6910	0.6842	0.6457	0.6378	0.6362	0.6320
overall Wald chi^2	85.72***	91.53***	94.31***	113.13***	105.16***	607.07***	546.11***
t on relative cohort	2.4**	2.5**	1.8*	2.8***	1.7*	4.6***	2.9***
t on highest Z	0.2	1.0	3.4***	1.2	1.7*	1.0	3.5****
t on highest Zy (Z * income/cap)	0.0	1.5	2.0*	1.8*	1.1	2.1**	5.6***
combined Wald chi^2 on all Z	na	na	na	na	na	30.82***	17.39***
combined Wald chi^2 on all Zy	na	na	na	na	na	48.45***	33.35***
Wald chi^2 on highest Z & Zy	na	na	na	na	4.95*	13.60***	46.06***
Wald chi^2 on highest Z	na	na	na	1.46	3.01*	1.02	12.00***
Wald chi^2 on highest Zy	na	na	na	3.22*	1.14	4.38**	31.30***
B. Equation (6): including percent foreign-born and percent urban in the basic model							
R-sq: within	0.8295	0.8293	0.8244	0.8111	0.7992	0.7909	0.7702
between	0.4300	0.4331	0.4436	0.4028	0.3976	0.3949	0.4130
overall	0.7522	0.7525	0.7504	0.7274	0.7187	0.7122	0.6996
overall Wald chi^2	151.27***	158.42***	164.21***	179.92***	164.13***	719.76***	650.90***
t on relative cohort size	1.9*	1.8*	1.5	2.9***	1.4	4.8***	3.3***
t on highest Z	1.4	1.2	7.9***	1.1	1.6	0.1	4.5***
t on highest Zy (Z * income/cap)	0.5	0.7	1.6	0.3	0.7	1.2	6.21***
combined Wald chi^2 on all Z	na	na	na	na	na	34.81***	20.97***
combined Wald chi^2 on all Zy	na	na	na	na	na	64.37***	48.77***
Wald chi^2 on highest Z & Zy	na	na	na	na	7.15***	15.83***	40.38***
Wald chi^2 on highest Z	na	na	na	1.18	2.59*	0.01	20.69***
Wald chi^2 on highest Zy	na	na	na	0.08	0.52	1.54	36.93***
C. Equation (6) plus growth rate interactions with age structure variables							
R-sq: within	0.8278	0.8282	0.8250	0.8110	0.7988	0.7916	0.7782
between	0.4495	0.4483	0.4546	0.4150	0.4132	0.4123	0.4324
overall	0.7543	0.7541	0.7524	0.7292	0.7209	0.7149	0.7082

Table A.2. (*Continued*)

Degree of polynomial:	**8**	**7**	**6**	**5**	**4**	**3**	**2**
overall Wald chi^2	160.13***	156.91***	164.92***	178.29***	163.58***	705.88***	673.23***
t on relative cohort size	1.9*	1.8*	1.5	2.9***	1.5	4.7***	3.9***
t on highest Z	1.1	0.7	2.9***	0.9	1.5	0.1	4.0***
t on highest Zy (Z * income/cap)	0.0	0.4	1.5	0.1	0.7	1.2	6.2***
combined Wald chi^2 on all Z	na	na	na	na	na	59.21***	48.89***
Wald chi^2 on highest Z & Zy	na	na	na	na	6.08**	9.06**	46.04***
Wald chi^2 on highest Z	na	na	na	0.80	2.19	0.01	16.33***
Wald chi^2 on highest Zy	na	na	na	0.01	0.44	1.46	38.56***
D. Equation (9): basic model plus % foreign-born, % urban, growth interactions, and year dummies							
R-sq: within	0.8437	0.8452	0.8445	0.8437	0.8429	0.8365	0.8363
between	0.7061	0.6949	0.6840	0.5697	0.4552	0.4694	0.4631
overall	0.8133	0.8126	0.8101	0.7895	0.7644	0.7605	0.7597
overall Wald chi^2	246.98***	260.83***	257.18***	278.68***	249.58***	924.45***	955.92***
t on relative cohort size	1.1	1.1	2.1**	2.2**	3.1***	2.3**	2.1**
t on highest Z	0.6	1.0	2.7***	2.1**	0.5	0.4	2.8***
t on highest Zy (Z * income/cap)	0.6	0.4	0.9	3.3***	1.5	0.6	6.8***
combined Wald chi^2 on all Z	na	na	na	na	na	8.77**	10.39***
combined Wald chi^2 on all Zy	na	na	na	na	na	46.19***	51.84***
Wald chi^2 on highest Z & Zy	na	na	na	na	5.69**	0.51	67.65***
Wald chi^2 on highest Z	na	na	na	4.49**	0.25	0.13	7.70***
Wald chi^2 on highest Zy	na	na	na	11.21***	0.14	0.34	45.87***

The standard test for the appropriate n starts from a high-degree polynomial ($n = 8$) and then successively adds restrictions until the last restriction fails. In the cases above, although t-statistics for the highest Z's are significant in a sixth-degree polynomial in Model A, and in a fifth-degree polynomial in model D (indicated in bold), Wald tests are singular in these cases, suggesting an over-fitted model. The only D supported by all test statistics is $n = 2$ (a quadratic polynomial, indicated in bold in all models).

Dependent variable is −ln (personal consumption expenditure/total personal income).

248 observations.

All variables except growth rate used in logged form.

Standard errors in parentheses. ***: $p < .01$; **: $p < .05$; *: $p < .10$.

A designation na indicates cases in which the Wald test is singular.

Estimated as random effects models using GLS to control for heteroscedasticity.

years of age 0–84 and 85+, J—the total number of age groups—is 86 in this analysis. The φ_j are constrained to sum to zero in order present comparable results across models.

However, it would be impractical to attempt to estimate 86 separate coefficients. As an alternative this analysis has adopted a method suggested by Fair and Dominguez (1991), which is in turn similar to Almon's (1965) distributed lag technique. The φ_j are constrained to sum to zero, and constrained to lie on a polynomial of degree n (where n is to be determined in fitting the model) such that

$$\varphi_j = \zeta_0 + \zeta_1 j + \zeta_2 j^2 + \zeta_3 j^3 + \cdots + \zeta_n j^n \qquad j = 1 \ldots J \tag{A.2}$$

and the following constraint has been imposed:

$$\sum_{j=1}^{J} \varphi_j = 0 \tag{A.3}$$

Substituting equation (3) into equation (2) produces

$$\zeta_0 = -\zeta_1(1/J)\sum_{j=1}^{J} j - \zeta_2(1/J)\sum_{j=1}^{J} j^2 - \cdots - \zeta_n(1/J)\sum_{j=1}^{J} j^n \tag{A.4}$$

and thus

$$\sum_{j=1}^{J} \varphi_j p_j = \zeta_1 Z_1 + \zeta_2 Z_2 + \cdots + \zeta_n Z_n \tag{A.5}$$

where

$$Z_n = \sum_{j=1}^{J} p_j j^n - \sum_{j=1}^{J} p_j \sum_{j=1}^{J} j^n \tag{A.6}$$

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