**Have We Put an End to Social Promotion?**

**Changes in Grade Retention Rates among Children Aged 6 to 17 from 1972 to 2003[*]**

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# Have We Put an End to Social Promotion?

## Changes in Grade Retention Rates among Children Aged 6 to 17 from 1972 to 2003

**Abstract**

We examine trends over time in age-grade retardation in schooling at ages 6 to 17 and in the effects of its demographic and socioeconomic correlates. We estimate a logistic regression model of age-grade retardation with partial interaction constraints using the annual October school enrollment supplements of the Current Population Survey. This model identifies systematic variation in the effects of social background across age and time from 1972 to 2003. While the effects of socio-economic background variables on progress through school become increasingly powerful as children grow older, that typical pattern has been attenuated across the past three decades by a steady, secular decline in the influence of those variables across all ages. A great deal of concern has been expressed about rising levels of economic and social inequality in the United States since the middle 1970s, and about the potential intergenerational effects of such inequality. However, there has been an opposite trend in the effects of social origins on age-grade retardation, which is an important indicators of progress through schooling. A trend is not a law, and there is reason to be concerned about the recent deceleration of the secular decline in effects of social background on age-grade retardation.

## INTRODUCTION

Grade retention is one of the methods often proposed and used to help poor performing students catch-up to their peers. At best, most research on the effects of grade retention portrays it as a practice that provides no benefit to the students; at worst it is considered a damaging practice. For example, Jimerson's (2001: 434) meta-analysis concludes that,

> "studies examining the efficacy of grade retention on academic achievement and socioemotional adjustment that have been published during the past decade report results that are consistent with the converging evidence and conclusions of research from earlier in the century that fail to demonstrate that grade retention provides greater benefits to students with academic or adjustment difficulties than does promotion to the next grade."

Jimerson found that retained students performed worse than their peers both academically and socially. Nagaoka and Roderick (2004) compare students close, both above and below, to the Chicago School System's test based retention cutoff score. They find that third graders who were retained showed no difference in test scores compared to those who were promoted and sixth graders who were retained continued to score worse than those who were promoted.

Previous studies have also shown that retained students are more likely to drop out of high school (Alexander, Entwisle, and Dauber 2003; Jimerson 2004; Rumberger and Larson 1998; Shepard 2004). Allensworth (2004) finds that the Chicago Public Schools' efforts to end social promotion did not change the overall dropout rate but that failing to pass the eighth grade promotional gate increased the likelihood that low achieving students would drop out. The most disturbing result was that the promotional gates policy exacerbated dropout differentials by race and gender; there was little change in the dropout rates of males and African Americans. Similarly, using national data from October Current Population Surveys, Hauser, Simmons, and Pager (2004) find that being over-age for grade is a powerful, proximate antecedent of high school dropout. Jimerson (1999) also reports that results from a 20 year prospective longitudinal study show that children who have been retained are less likely to earn a diploma by age 20, less likely to attend post-secondary education, were paid less, and have lower employment ratings than a group of promoted low achieving students.

Beebe-Frankenburger et al. (2004) call the entire practice of retention into question. They find evidence that compared with their at-risk and promoted peers, retained students are not significantly younger and have lower IQ scores. They argue that these findings contradict the assumptions underpinning the rationale for retention; that it benefits children who are immature and that retained children have the capability to meet standards during the extra year. If one takes age as a rough proxy for maturity, as the authors do, then the children who are retained in practice will not benefit from it because they are not significantly younger. Additionally, with their lower IQ scores, the retained children may not be able to meet the standards the second time through the material. Shepard (2004) echoes this sentiment, arguing that based on the research literature to date, retention does not meet the Food and Drug Administration's guidelines of a safe and effective treatment as used in the drug approval process.

Some recent studies, not included in Jimerson's (2001) meta-analysis, portray a more favorable view of the effects of retention. Alexander, et al. (2003) find positive results of retention on academic achievement in a longitudinal study of Baltimore school children. Their favorable finding is that retention halts the downward slide that retainees were experiencing previously in their academic careers. These positive results only appear for children retained only once in elementary school and after the first grade. However, Alexander, et al. acknowledge that, consistent with other studies, these effects diminish over time. In their critique of the first edition, Shepard, Smith and Marion (1998) note that in the multivariate analyses, the positive effects of retention wash out.

Shepard (2004) raises the following critiques of the analysis. The authors misinterpret the greater test score gains of retained students relative to their peers and fail to fully consider regression to the mean of the test scores or effects of repeated testing. They also overemphasize the beneficial effects experienced by second and third grade retainees while first grade retainees show detrimental effects of retention. Hauser (2005) notes that the analyses performed by Alexander, et al. are circular in nature; they control for events that are subsequent to the retention decision. He adds that effects of race and gender are not reported in the book, and the descriptive statistics reported in the book are wrong because they are unweighted in spite of the oversampling of white, higher status students in the Baltimore study. Moreover, like other studies, Alexander, et al. (2003) find much higher rates of school dropout among students who had been retained.

Eide and Showalter (2001) use an instrumental variable approach[1] with data from the High School and Beyond (HSB) and the National Education Longitudinal Study (NELS:88). Their study attempts to isolate the effects of retention on high school completion and labor force earnings while avoiding possible simultaneity between retention and academic outcomes. They

[1] They use the difference in days between the cutoff date of kindergarten entry and the child's birthday as their instrumental variable.

give a cautiously optimistic reading of findings that indicate beneficial effects of retention that are not statistically significant. Likewise, Jacob and Lefgren (2004) use an instrumental variable approach with administrative data from the Chicago Public School system and find that for those students within a narrow range of the cutoff test score determining retention and summer school, there are modest positive effects on achievement among white third graders and no effects on 6th graders who have been retained. Consistent with other studies, Jacob and Lefgren's findings indicate that whatever advantage exists for retained students is temporary; the effects on test scores two years after the repeated grade are smaller than those one year after retention.

Despite these negative or weak findings, the popular sentiment in America is that schools and teachers need to be more accountable to ensure that children progress at appropriate rates (National Education Goals Panel, Goal #8, The National Commission on Excellence in Education, 1983). Many politicians, including Presidents Clinton and Bush, have made direct pleas to end social promotion (Bush 2004; Clinton 1999). Hauser (2004, 2005) warns that the annual testing in third through eighth grades mandated by the No Child Left Behind Act (US Congress 2002) may increase the incidence of test-based retention. That is, whatever the purposes, strengths, or weaknesses of tests, once given they tend to be used to make decisions about test-takers.

There is some agreement that the rate of children who have ever been retained is growing (Allington and McGill-Franzen 1992 for New York State; Hauser, Pager, and Simmons 2004; Jimerson and Kaufman 2003; McCoy and Reynolds 1999; Shepard and Smith[2] 1989 for 14 states and Washington, D.C.) but it may not affect all subpopulations equally. Indeed retained students tend to be male (Corman 2003; Hauser, Pager, Simmons 2004; Zill, Loomis, and West 1997) ,

[2] Shepard and Smith (1989: Table 1 p 6)

black or Hispanic (Alexander, et al. 2003; Hauser, Pager, Simmons 2004; Jimerson, et al. 1997; Zill, et al. 1997), younger relative to their peers (Corman 2003; Shepard and Smith 1986) and from disadvantaged backgrounds, e.g., from broken families or with lower parental educational attainment and lower family income (Corman 2003; Hauser 2001; Hauser, Pager, Simmons 2004).

Has the pressure to end social promotion translated into more children being retained? A definitive answer to this question is not easy to come up with, as the intense debate implies – and the full testing regime mandated by the No Child Left Behind Act is not yet in place. Shepard (2004) notes that retention and social promotion are not mutually exclusive. The prevalence of policies limiting double retentions means that a student may be retained in one year and socially promoted in following years. Corman (2003) estimates retention rates using the National Household Education Survey (NHES). He finds that, between 1991 and 1996, 10 to 15% of six year olds and 28 to 30% of fifteen year olds have been retained. Wheelock (2005) reports estimates from NELS:88 that 20% of eighth graders have been retained, and 21% of adolescents in the ADD Health Study have been retained.

The first place one would think to look for information about grade retention is the National Center for Education Statistics (NCES). While the *Digest of Education Statistics 2003* does not mention social promotion or grade retention, the NCES data have improved since 2002 regarding grade retention. *The Condition of Education in 2005* discusses delayed entry to and retention in kindergarten. It compares differentials between on-time kindergarteners, delayed-entry kindergarteners and kindergarten repeaters from the Early Childhood Longitudinal Study – Kindergarten Class of 1998 (ECLS-K). However, there is no report of the overall prevalence of even this single measure of retention and no mention of retention in elementary or secondary school at all.

Hauser (2001, 2004) mentions three possible sources of national-level data where one might find national trends in grade retention over time. The first source of retention information is exemplified by the state data collected by Shepard and Smith (1989) and, more recently, by the National Research Council (Heubert and Hauser 1999).[3] These data have the advantage of being direct reports from state education agencies of the incidence of grade retention. The disadvantage is exemplified by the NRC report: the average number of years in which the 26 states plus Washington DC provide estimates of retention is 4.22 years, and most of these years occur in the mid-1990s. Thus, the data are no longer timely, and geographic coverage is limited. Shepard and Smith (1989) note that not much about the comparability of state retention data is known, specifically whether the population used in the denominator is from beginning or end of the year enrollment data.

Retention estimates have also been made using the Census (Hauser 2004) and the October Supplement to the Current Population Survey (CPS) (Bianchi 1984; Eide and Showalter 2001; Hauser 2001, 2004; Hauser et al. 2004). These estimates yield the longest time span but the drawback is that retention has to be inferred from age and current grade enrollment in the CPS or educational attainment by age in historical census data.[4] Although it lacks a repeated

[3] *High Stakes Testing* table 6-1.

[4] Eide and Showalter (2001) compare the grade enrolled in last year with the grade enrolled in this year. These data however are only available in the CPS in and after 1994. This is a qualitatively different measure because our measure is a proxy for ever being retained while their measure only captures retention in the current year of the survey.

measure of grade retention or measures of academic achievement, the CPS is the best data with which to construct comparable yearly estimates of the incidence of grade retention on a national level (Hauser 2001).

In order to estimate the retention rates in the CPS data, we constructed a measure from the age and currently enrolled grade of each student between the ages of 6 and 17. This measure identifies children who are below the modal grade for their age and has been used to estimate retention rates in prior studies (see Hauser, Pager, Simmons 2004; Corman 2003 uses a similar measure with month and year and state cutoff dates to corroborate retrospective reports).

**INCIDENCE OF RETENTION**

Figure 1 reports the proportion of children at selected ages who are enrolled below the modal grade for their age (hereafter, BMG). In order to smooth the lines, the trend lines are three year moving averages and the horizontal axis indicates the year in which the cohort was born. Read vertically, it shows the within-cohort change in the incidence of BMG. As expected, the proportion of BMG children increases at each age except for a crossover between ages 15 and 17 for the 1983 through 1985 cohorts. Through the 1986 cohort, the last year for which we have complete cohort data, the later cohorts have higher overall incidence levels than the earlier ones. The increase in the incidence of BMG occurred within a few years. On each side of this increase the rates were relatively stable. Cohorts born after 1986 seem to show a decrease in the proportion of children who are behind their peers at ages six and nine.

Does this trend reflect actual changes in the retention rates or is it merely an artifact of increasing age at entry into school or other non-retention phenomena? The lines are very nearly parallel to the trend line for six year olds suggesting that the overall BMG rates are driven by age at school entry and/or by retention at early ages. There is a direct retrospective question about grade retention in the school enrollment section of the CPS but unfortunately it is only asked in 1992, 1995, and 1999. There is yet a third way to construct estimates of retention trends with CPS data. After 1994, respondents were asked not only about the current grade but also the grade in which students were enrolled in the previous year (Eide and Showalter 2001). This measure is different from the other two because it only picks up retentions in the last year, not ever having been retained as the other two measures do.

Table 1 compares the constructed BMG measure with the direct retrospective report of whether the student has ever been retained. If we assume the latter is the true measure, more than eighty percent of children are correctly classified by BMG. The overwhelming majority of misclassified children are false positives (being BMG without having been retained). Why would the BMG measure err in the direction of false positives? Among other things,[5] it cannot distinguish between children who had been retained in grade and those whose parents decided to delay school entry. Hauser and colleagues (Hauser 2004; Hauser, Pager, Simmons 2004, Heubert and Hauser 1999) have used the school enrollment supplements to the CPS to document the rates of children who have been retained between ages 6 and 17. They used 6 year-olds who were below the modal grade for their age as a baseline measure of BMG for that cohort in order to isolate retentions that happened after first grade. By subtracting this cohort baseline for each of the successive age groups, they are able to identify retention that happens after entry into elementary school. Figure 2 updates their series. The horizontal axis is the birth year of the cohort. The trend lines are 3 year moving averages in order to smooth fluctuations.

[5] Other reasons could be children whose birthday occurs between state mandated cutoff dates and the administration of the survey (see Cascio 2005) or children who have missed a year of school due to health reasons.

There is complete cohort data through the 1986 cohort. The BMG rates after age 6 were fairly steady with 8 to 10% of children BMG by age 8, 13 to 17% by age 11, 17 to 21% by age 14, and 22 to 24% by age 17 through the cohort of 1978. Then there was a sharp decline in BMG at all ages to 7%, 11%, 14% and 17% for the four age groups respectively. The rates have seemed to stabilize around these values since then. The most recent data from the two youngest age groups indicate that there might be another upswing in retention during the elementary school grades. Figure 2 corroborates the evidence from figure 1 that the rise in retention may be driven by pre-first grade retentions or delays in school entry. It appears that retentions at older ages are being replaced by retentions at younger ages.

The rise in 6 year olds who are BMG from 1978 to 1992 (the cohorts of 1972 to 1986) coincides with Shepard's 1989 observation that, "Holding children back in kindergarten in large numbers is a phenomenon of the 1980s."[6] She concludes that retention in kindergarten is still retention; some of the increase in BMG 6 year olds is attributable to retention per se. Another reason for the increase is "academic red-shirting." Marshall (2003) reviews research on academic red-shirting. Whether the decision to red-shirt children is made by their parents alone or with teacher input, the reasons given are similar to those given for retaining children. Either the child needs more time to mature, or the extra year would give a lower performing student a chance to catch up to meet the expectations of kindergarten. The effects of red-shirting reported by Marshall are similar to effects of retention in that there is a temporary advantage to the red-shirted child, but it disappears by third grade. Red-shirted children are also more likely to be placed in special education classes than comparable peers. Because of the similarity of red-

[6]Shepard, 1989:65

shirting and retention, the number of false positives identified among BMG children may not be as great as it appears.

Cascio (2005) estimates the bias introduced by using BMG as a proxy for grade retention in regression estimates of the effects of social background. Her sample includes children aged 7 to 15 in 1992, 1995, and 1999. She finds that when used as a dependent variable, the effect of the misclassification of students in the CPS data set attenuates the magnitude of the true coefficients of the independent variables by as much as 35 percent. The attenuation bias is greater for males (38% vs. 32%) and for older ages - 41% for ages 12 to 15 versus 30% for ages 7 to 11. The attenuation bias decreased over time from 40% in 1992 to 32% in 1999. The attenuation bias is highest for non-Hispanic blacks (45%) followed by non-Hispanic whites (34%), Hispanics (31%) and the other races are grouped together and have an attenuation factor of 28%.

**DATA**

The CPS data on school enrollment come from a nationally representative probability sample of the civilian, non-institutionalized population each October. While they do not have detailed educational measures, such as the previous year's enrollment (prior to 1994) or information about transitional or special education classrooms, the data do provide repeated cross sections of the national population over a long period of time. The fact that the data are repeated cross-sections, rather than true longitudinal observations, does not pose a problem because we are interested in aggregate retention rates, rather than in the consequences of retention for individuals. We are able to construct a uniform CPS file from 1968 onward because a common set of social background questions have been asked every year, along with information about age and grade enrollment of school-aged persons (Hauser & Hauser 1993; Hauser, Jordan, & Dixon 1993); however, prior to 1972, the data do not capture Hispanic ethnicity.

Table 2 lists the two sets of covariates used in the analysis: demographic and socioeconomic variables. The demographic covariates include gender, race, age, year of participation in the survey, region of the country, metropolitan status and number of siblings. The socioeconomic variables include the education and occupational status of the household head and his or her spouse, family income in the twelve months immediately prior to the survey, [7] whether or not the child's family owns their home, and whether the child comes from a broken family.[8] The education measures varied over time. Through 1991, the variable was measured as years of education from no years of education to 6 years of college or more. From 1992 forward, education was measured as highest category of school or degree completed. Hauser (1997) discusses the incompatibilities and inconsistencies of these two measures of education. For the analysis presented here, we converted the educational credentials to the metric of putative years of school completed. This introduces a degree of measurement error by assuming, for example, that 16 years of education is identical to a bachelor's degree.

There are two education variables for each 'parent' in the final model. The first captures the number of years completed through high school graduation, and the second captures college education. Thus, a person who completed the 10$^{th}$ grade has a value of 10 for the first variable

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[7] Family income is collected in ranges in the October CPS that have changed over the years. The variable is the natural log of the midpoints of these ranges adjusted to constant dollars with the CPI-U series published by the BLS.

[8] This is defined as not living with a mother and father (female head with male spouse or male head with female spouse).

and 0 on the second. The corresponding scores for high school graduates with no college and people with a bachelor's degree are 12, 0 and 12, 4, respectively. This scheme was used to allow for piecewise linear effects of education before and after the high school to college transition, especially in light of the differences in college attendance over time among the parents in this sample.

The CPS data include 864,878 children aged 6 to 17 clustered within 256,608 households over 32 years. In the multivariate analyses below, we report results from a one half random sub-sample stratified by age, period, gender and race. This sub-sample contained 432,626 observations clustered within 199,278 households. All of the models reported here were estimated using robust standard errors that have been adjusted for household clustering. There were missing data in some observations on seven continuous variables. The fourth and fifth columns of Table 2 show the differences between the data without missing data and the data with imputations. The final column of Table 2 shows the total number of non-missing observations for each variable. Appendix A shows the percent missing on each variable by year.

There are two types of missing data on the variables used in this analysis. Data on household income and head's education are sometimes missing but there are real values associated with these characteristics out in the world.[9] We used multiple imputation to replace this type of missing data. When data are not missing completely at random (MCAR), listwise deletion can yield biased results (Allison 2001:6). Multiple imputation yields consistent, asymptotically efficient, and asymptotically normal estimates under the weaker assumptions that the data are missing at random (MAR) and the model is correctly specified (Allison 2001). Of

[9] A third type of missing data is present but is imputed using a hot deck method by the Census Bureau. For the sake of this analysis, we treat these imputations as unproblematic.

the two variables with this type of missing data, income is the only one for which there is evidence that it is not MAR. However, only seven percent of cases lack income data, so this violation should not strongly affect the estimates.

The data on spouse's education and occupational status as well as head's occupational status are missing for a different reason; values do not exist because of unemployment or because there is no spouse in the household. In order to account for this type of missing data we used a dummy variable adjustment technique. We substitute the missing data on these variables with mean values of non-missing cases, conditional on the age of their child and the period of the survey. Then we included a dummy variable in each model that we estimated, indicating whether the observation had missing values. Where missing observations truly do not exist, this method of accounting for missing data is statistically sound (Allison 2001; King, et al. 2001).

## MODEL

There are two prototypical ways to estimate trends over time and age using conventional logistic regression (logit) models. Equation 1 is a logit model with BMG as the response variable, where *i* indexes individuals, *j* indexes age, and *k* indexes time periods. The social background variables are indexed by *l*. The 32 years are divided into eight groups of four years that coincide with presidential terms. In order to accomplish this division, the first period covers five years, and the last period covers three years. Thus, the alphas represent separate intercepts for each age by period combination, the variables of interest in this simple model. This model is unsatisfactory because the only evidence of change is in the different levels of the intercepts; the effect sizes do not change.

$$\text{logit}[P(Y_{ijk} = 1)] = \alpha_{jk} + \sum_{l} \beta_l x_{il} \tag{1}$$

In a more nuanced treatment, Hauser, Pager, Simmons (2004) used separate logistic regressions at selected ages to examine the effects of demographic and social background variables as well as time on being BMG. This allowed them to look at how the effects differed among children aged 6, 9, 12, 15, and 17. A similar set of logistic regressions is required to assess differences in how the effects of social background variables change with age and time. Equation 2 is the full interaction model in which the social background effects can vary independently in each age by period category. Again, the alpha term represents differential intercepts. Now the beta term represents the effects of social background variables at the baseline category, age = 6 and period = 1972 – 1978. The gamma term contains the differences in effects over age and period. In each of these models, the effects vary independently over each combination of age and period. This model allows all the effect sizes to vary unconstrained but uses many more of degrees of freedom, especially as the number of categories of *j* and *k* gets bigger.

$$\text{logit}[P(Y_{ijk} = 1)] = \alpha_{jk} + \sum_{l} \beta_l x_{il} + \sum_{l} \gamma_{jkl} x_{il} \qquad (2)$$

An equivalent method is to estimate the logit model in equation 1 for each age-period combination. Either way, this model yields unwieldy results because there are more than 2000 estimated parameters.

Equation 3 is a logit model with interaction constraints (LIC), the main model of interest in this paper.[10] The alpha term represents an intercept for each level of age, indexed by *j*, and time period, indexed by *k*. The next term consists of a vector of parameters, $\beta$, for the vector of

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[10] The model in equation 3 is easily estimated with standard statistical software. Appendix B contains an example of the Stata code used in the estimation.

explanatory variables, $X_{kl}$ . The final two terms, containing $\lambda_j$ and $\lambda_k$, capture the proportional change in the linear predictor separately for each level of *j* and *k*. The proportional change, $\lambda_j$, is the factor by which the linear predictor changes as age increases, while $\lambda_k$ is the factor by which the linear predictor changes for each time period.

$$\text{logit}[P(Y_{ijk} = 1)] = \alpha_{jk} + \sum_l \beta_l x_{il} + \lambda_j (\sum_l \beta_l x_{il}) + \lambda_k (\sum_l \beta_l x_{il}) \tag{3}$$

The advantage of this model is that it is more detailed than the model that constrains change in the effects to be identical across ages and time, but more parsimonious than the full interaction model in which effects may vary independently by age and by time.

If $\lambda$ is positive, it means that the effects of the explanatory variables increase in magnitude. Similarly if $\lambda$ is negative, the effects decrease in magnitude. This can be shown by factoring the linear predictor out of the final three terms:

$$\text{logit}[P(Y_{ijk} = 1)] = \alpha_{jk} + (1 + \lambda_j + \lambda_k) \sum_l \beta_l x_{il} \tag{4}$$

The $\lambda$ variables can either be continuous, ordinal, or nominal. In order to test for model fit, we ran four different models in which we treated both age and period as continuous variables, both as nominal variables, and one as nominal while the other was continuous. When the lambda variable is continuous, $\lambda_5$ is half the size of $\lambda_{10}$. When the lambda is nominal, the interpretation of $\lambda_k$ is relative to an omitted baseline category. We only report results from the models where age and period are nominal because they fit the data best. Also, Morris (1993) found that retention rates across grade levels were not linear but were better described by a negative growth exponential function.

Previous empirical work with the data showed that the unconstrained model is preferred to the LIC model in this case.[11] Our results did show that there was a group of variables for which the constraints worked well. We decided to relax the constraints on the ill fitting covariates and came up with a hybrid of the models in equations 2 and 4. We call this model (equation 5) a logit with partial interaction constraints (LPIC). In

$$\text{logit}[P(Y_{ijk}=1)] = \alpha_{jk} + (1+\lambda_j+\lambda_k)\sum_{1}^{l}\beta_l x_{il} + \sum_{l+1}^{L}\gamma_{jkl}x_{il} \tag{5}$$

this equation, the explanatory variables are separated into two groups. There are $l$ explanatory variables with constrained interactions. The remainder of the explanatory variables, from $l+1$ to $L$, have unconstrained interactions. We divided the variables *post hoc* according to previous empirical work mentioned above. The seven constrained variables are: gender, income, home ownership, head's K-12 education, head's occupational status, spouse's occupational status, and the total number of children in the household.

**MODEL FIT**

Table 3 lists the F-statistics of the model comparison test for multiply imputed data proposed by Allison (2001:68). Both the LPIC model and the unconstrained model are preferred over the simple logistic regression. The last row of Table 3 tests the hypothesis that the LPIC model fits as well as the unconstrained model. The p-value of 0.25 in the last column indicates that the unconstrained model does not significantly improve the fit over the LPIC model at conventional levels despite the statistical power of over 430,000 cases. Based on this statistical evidence and its greater parsimony, we will discuss the results from the LPIC model.

**FINDINGS**

**Overall Odds of Retention**

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[11] For a thorough discussion of the decision process see Frederick, 2005.

Table 4 lists the predicted log odds of being BMG for the average child in the sample, i.e. evaluated at the means of the covariates. This illustrates how the log odds of being BMG have changed across ages and time periods. The surface plot in Figure 3 indicates that the highest predicted log odds of being BMG occur at the oldest ages during the most recent periods. Net of social background characteristics, the log odds of being BMG have increased as children age in each time period, supporting the common sense hypothesis that the longer children are exposed to the risk of being retained, the more they are actually retained. Figure 4a, further shows that these increases are largely parallel and monotonic in every time period. The biggest increases occur between the ages of 6 and 8, flatten out until age 14 and then increase slightly at ages 15 and beyond.

Figure 4b shows the change in log odds of being BMG over time for each age group. All age groups show an increase in the log odds of being BMG over time. For children aged 6 and 7, there is a sharp increase from 1972-1976 to 1989-1992 and the trend levels off or slightly declines after that. At older ages, the increase in the log odds during the early periods is less severe than those for 6- and 7-year-olds because the trend lines start at higher levels. The maximum predicted values occur during the 1989-1992 or 1993-1996 periods and stabilize until the current periods. This evidence provides further support to Shepard's characterization that the 1980s were the decade of kindergarten retention.[12] It also suggests that retentions in the first few grades of elementary school are being replaced by either retentions in kindergarten or first grade or academic red-shirting by the 1990s because the gap between 6 and 7 year olds declines rapidly.

**Constrained Variables**

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[12] Shepard. 1989. "A Review of Research on Kindergarten Retention."

The coefficients on the constrained covariates are all in the expected directions. Boys are more likely than girls to be BMG. Each additional child in the household increases the chances of being BMG. Children in families who own their homes, have higher household income, in which the household head has more education and both the head and the spouse have higher occupational status scores all are less likely to be BMG.

Alone, the coefficients of these variables represent their effects on children who were 6 years old between 1972 and 1976. Figure 5a displays the age interaction constraint estimates. These values are all positive and increase linearly except for a dip at age thirteen. Net of the other covariates and changes in period, as children age the coefficient increases by a factor of $(1+\lambda_j)$. For example the coefficient for boys in the first period is 0.32, 0.48, 0.58, 0.62, and 0.69 at ages 6, 9, 12, 15, and 17 respectively. For each of the seven constrained covariates, the magnitude more than doubles from the time children enter school to the time they finish. These aspects of social background increase in importance as children age.

Figure 5b presents the estimates of the period interaction constraints. Because these constraint factors are all greater than negative one, the quantity $(1+\lambda_k)$ approaches zero, thereby reducing the magnitude of the coefficients. More concretely, the coefficient of logged household income for six year olds approaches zero from -0.13 in 1972-1976 to -0.12 in 1985-1988 to -0.06 in 2001-2003.[13] Independent in the changes in age, there has been a secular decline in the importance of these seven social background characteristics in predicting being BMG over time.

This decrease is not as linear as the trend in age. Instead, this trend resembles four steps. The first two drop-offs occur between the second and third and between the fourth and fifth

[13] The odds ratios are not shown here but can be calculated using the interaction constraint factor implied in equation 4 and the estimates in Table 5.

periods. A third steeper drop-off takes place during the sixth and seventh period. Evidence from the latest period might signal a reversal of this trend because the interaction constraint for 2001-2003 is greater than that of 1997-2000. This difference is not statistically significant and only time will tell whether trend toward decreasing effects of social background has ended.

The estimates from our LPIC model indicate that net of everything else the differences between the group of children who are BMG versus those that are progressing through school at a normal rate are becoming increasingly similar with respect to these seven characteristics. Quick examination of this evidence could lead to the naïve conclusion that retention decisions are increasingly made on a class and gender blind basis. One of the only things left that could differentiate these populations that is not controlled for in this model is the ability of the student. This conclusion falters, however, if one keeps in mind that real achievement gaps persist along gender and class lines. If children are retained due to poor achievement and achievement is correlated with gender and class we should expect to see a correlation between retention and gender and class. This line of reasoning renders our results paradoxical. One possibility is that in the collective rush to end social promotion, schools are retaining greater numbers of children in a more capricious manner than ever.

**Unconstrained Variables**

No such unifying trend was found for the rest of the covariates. In the rest of this section we discuss three groups of these covariates: race, region and city status.[14] The only statistically significant racial difference was between blacks and whites. Table 6 lists the odds ratios of being BMG for black children relative to non-Hispanic whites. At each age, the effects stay relatively constant over time. Black children are generally less likely to be BMG than whites

[14] The full results are available upon request from the authors.

before age 9. After age 9, black children are more likely to be BMG. This gives a more nuanced view of the common finding that black children are more likely to be retained that whites. It does not look like it is as simple as blacks being retained more than whites at each level of schooling. Instead, our evidence suggests that despite the fact that black children are more likely than whites to start school on time; disproportionately more black children are being held back at older ages, resulting in a shift the odds in their (dis)favor. Future work with more direct measures including the timing of retention is required to corroborate this finding.

The differences between the Northeast and the Midwest, South and West follow distinct patterns. The differences between the Midwest and Northeast generally decline so that at young ages, Midwestern children are much more likely to be BMG but at older ages, the differences are close to zero. The trend over time at each age level takes on more or less a U shape. Midwestern children are more likely than Northeastern children to be BMG at the beginning and end of the time series. Comparing the South to the Northeast, there is no discernable trend in the changes in magnitude as children age. The most consistent change over time at most ages is that there is an increase in the likelihood of being BMG between the two most recent periods. The differences between the West and the Northeast bounce around above and below zero. There is a small general decline in the difference as children age. The most striking trend is that, at each age the magnitudes of the coefficients converge around their lowest point in the 1993 to 1996 period.

The differences between children in major central cities and children in other places follow the same general trend. The differences over age generally decline in each time period. The exception is in the current period in which there is, if anything, a slight increase for older children. At older ages, the differences between the two groups increase slightly. At younger ages, the differences increase to a high point in the second to fourth time period and then decrease after that.

**CONCLUSION**

After controlling demographic and socioeconomic background variables, the prevalence of children being below the modal grade for their age has increased overall from 1972 to 2003. This is a serious matter because age-grade retardation is a major, proximate cause of high school dropout. While our data do not allow us to tie actual retention decisions to public policy, it seems likely that the increased pressure for schools to be accountable and use of standardized tests will continue the upward aggregate trend in age-grade retardation. Regardless of the degree to which retention decisions are made based upon meritocratic ideals, increasing the proportion of the population without at least a high school diploma is a matter we should seriously evaluate.

What are the differences between children who are and are not below the modal grade for their age? Net of the changes over time, the seven constrained variables show increasing differences as age increases. Part of this trend is attributable to the fact that, ignoring the practice of academic red-shirting, unlike comparisons of test scores, almost all children enter school at roughly the same age with no retentions. Differences in retention by social background characteristics should only increase as children age, insofar as retention is correlated with these characteristics. This evidence indicates that there is a cumulative disadvantage associated with being male and having lower socioeconomic characteristics.

The LPIC model yields diminishing effect sizes over time for the constrained variables. What does this tell us about social promotion? The retained and never retained groups are increasingly more similar in social background, especially after 1988. This result contradicts evidence about achievement gaps along gender and class lines that would not predict such a convergence. The latest results from the NAEP long term trend data[15] show that the male-female gap in reading only converges among 9 year olds and the gender gap in math only converges for 17 year olds. The differences based on parental education have not appreciable changed since 1978. We think that it will be informative to continue to follow trends and differentials, not only in age-grade retardation but also in annual retention rates by grade as the effects of No Child Left Behind come to dominate educational outcomes in primary and secondary schools.

Because the LPIC model does not constrain all variables, we are able to observe a different pattern of differences between blacks and whites. These differences are relatively stable over time and show that Black children are at an extreme disadvantage regarding the probability of retention. This evidence is consistent with retention decisions based on academic achievement, reflecting the persistent gap between blacks and whites.

[15] Available on the web at http://nces.ed.gov/nationsreportcard/ltt/results2004/ (Last Accessed on 02/27/06)

## APPENDIX A

| The percentage of missing values on each variable by year. | | | | | |
|---|---|---|---|---|---|
| Year | Income | Head's Education[a] | Head's Occupational Status | Spouse's Education[a] | Spouse's Occupational Status |
| 1972 | 6.00 | 0 | 10.15 | 18.26 | 60.52 |
| 1973 | 6.95 | 0 | 9.87 | 18.41 | 58.94 |
| 1974 | 6.72 | 0.03 | 10.92 | 19.70 | 58.54 |
| 1975 | 7.50 | 0 | 11.23 | 20.60 | 58.08 |
| 1976 | 7.86 | 0 | 10.74 | 20.32 | 56.77 |
| 1977 | 7.78 | 0 | 10.46 | 20.87 | 55.09 |
| 1978 | 7.95 | 0 | 10.55 | 22.16 | 53.57 |
| 1979 | 7.46 | 0 | 9.72 | 22.25 | 52.68 |
| 1980 | 5.91 | 0 | 10.74 | 23.17 | 53.06 |
| 1981 | 4.78 | 0 | 10.82 | 23.96 | 52.47 |
| 1982 | 4.97 | 0 | 10.88 | 24.49 | 51.87 |
| 1983 | 4.30 | 0 | 10.95 | 24.92 | 52.45 |
| 1984 | 4.38 | 0 | 11.10 | 25.18 | 51.06 |
| 1985 | 3.56 | 0 | 11.09 | 25.36 | 49.61 |
| 1986 | 2.36 | 0 | 10.41 | 26.00 | 49.15 |
| 1987 | 3.43 | 0 | 10.79 | 26.44 | 48.94 |
| 1988 | 4.88 | 0 | 10.99 | 26.84 | 48.04 |
| 1989 | 6.40 | 0 | 10.98 | 26.89 | 47.27 |
| 1990 | 6.60 | 0 | 11.01 | 26.93 | 47.61 |
| 1991 | 5.22 | 0.13 | 12.49 | 27.78 | 48.26 |
| 1992 | 5.33 | 0.14 | 12.44 | 28.31 | 48.05 |
| 1993 | 5.68 | 0.18 | 11.81 | 28.17 | 47.44 |
| 1994 | 6.93 | 0.08 | 13.20 | 28.51 | 47.50 |
| 1995 | 8.64 | 0.26 | 13.16 | 28.38 | 47.18 |
| 1996 | 8.57 | 0.17 | 11.91 | 28.52 | 46.52 |
| 1997 | 8.85 | 0.17 | 11.49 | 28.72 | 47.47 |
| 1998 | 9.57 | 0.33 | 10.83 | 29.27 | 47.82 |
| 1999 | 10.82 | 0.40 | 11.18 | 29.68 | 48.47 |
| 2000 | 12.66 | 0.33 | 11.09 | 30.14 | 49.02 |
| 2001 | 12.21 | 0.35 | 10.89 | 30.22 | 48.43 |
| 2002 | 12.31 | 0.39 | 10.97 | 30.75 | 49.68 |
| 2003 | 14.70 | 0.48 | 11.48 | 30.27 | 49.79 |

[a] The two variables for both head and spouse's education are constructed from one variable so they are missing on the same cases

**APPENDIX B: Sample Stata code to estimate the LIC model**

*Each period represents a new line in the program*

```
.ml model lf lic (alpha: bmg = age7 age8 age9 age10 age11 age12
     age13 age14 age15 age16 age17 yr7780 yr8184 yr8588 yr8992
     yr9396 yr9700 yr0103) (beta: black hispanic othrace male
     majsub othercc othersub other midwest south west lninc
     ownhome brkfam headed headcolge spsed spscolge hsei ssei
     totkids, nocons) (lambda1: age7, nocons) (lambda2: age8,
     nocons) (lambda3: age9, nocons) (lambda4: age10, nocons)
     (lambda5: age11, nocons) (lambda6: age12, nocons) (lambda7:
     age13, nocons) (lambda8: age14, nocons) (lambda9: age15,
     nocons) (lambda10: age16, nocons) (lambda11: age17, nocons)
     (lambda12: yr7780, nocons) (lambda13: yr8184, nocons)
     (lambda14: yr8588, nocons) (lambda15: yr8992, nocons)
     (lambda16: yr9396, nocons) (lambda17: yr9700, nocons)
     (lambda18: yr0103, nocons), maximize difficult robust
     cluster(hhid)
.ml display
```

*The first command calls the following program, slm_ml18.ado, to do the estimation*

```
.program define lic
.args lnf alpha beta lambda1 lambda2 lambda3 lambda4 lambda5
     lambda6 lambda7 lambda8 lambda9 lambda10 lambda11 lambda12
     lambda13 lambda14 lambda15 lambda16 lambda17 lambda18
```

```
.tempname theta

.gen double `theta' = `alpha' + `beta' + (`lambda1'*`beta') +

     (`lambda2'*`beta') + (`lambda3'*`beta') +

     (`lambda4'*`beta') + (`lambda5'*`beta') +

     (`lambda6'*`beta') + (`lambda7'*`beta') +

     (`lambda8'*`beta') + (`lambda9'*`beta') +

     (`lambda10'*`beta') + (`lambda11'*`beta') +

     (`lambda12'*`beta') + (`lambda13'*`beta') +

     (`lambda14'*`beta') + (`lambda15'*`beta') +

     (`lambda16'*`beta') + (`lambda17'*`beta') +

     (`lambda18'*`beta')

.quietly replace `lnf' = ln(exp(`theta')/(1+exp(`theta'))) if

          $ML_y1==1

.quietly replace `lnf' = ln(1/(1+exp(`theta'))) if $ML_y1==0

.end
```

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| Table 1: Comparison of Retention Measures in the CPS | | |
|---|---|---|
| BMG | Direct Measure of Repetition | |
| | Yes | No |
| Yes | 5,768 | 12,418 |
| No | 822 | 49,792 |

| Table 2 - Covariate Proportions, Means and Standard Deviations for Complete Case and Imputed Datasets | | | | | |
|---|---|---|---|---|---|
| | Complete Case | | Imputed | | |
| Variable | Proportion | SD | | | No. Obs. |
| White | .7397 | 0.44 | | | 864878 |
| Black | .1340 | 0.34 | | | 864878 |
| Hispanic | .0855 | 0.28 | | | 864878 |
| Other Race | .0409 | 0.20 | | | 864878 |
| Male | .5103 | 0.50 | | | 864878 |
| Major Central City | .0900 | 0.29 | | | 864878 |
| Major Suburb | .1338 | 0.34 | | | 864878 |
| Other Central City | .1312 | 0.34 | | | 864878 |
| Other Suburb | .2172 | 0.41 | | | 864878 |
| Rural | .4278 | 0.49 | | | 864878 |
| East | .2135 | 0.41 | | | 864878 |
| Midwest | .2588 | 0.44 | | | 864878 |
| South | .2992 | 0.46 | | | 864878 |
| West | .2285 | 0.42 | | | 864878 |
| Home Ownership | .7180 | 0.45 | | | 864878 |
| Broken Family | .2172 | 0.41 | | | 864878 |
| | Mean | SD | Mean | SD | |
| Total Children in the Household | 2.73 | 1.45 | | | 864878 |
| Logged Income | 9.91 | 0.85 | 0.30 | 0.46 | 803533 |
| Head's K-12 Education | 11.10 | 2.07 | 11.10 | 2.07 | 864054 |
| Heads Postsecondary Education | 1.27 | 1.96 | 1.27 | 1.96 | 864054 |
| Spouse's K-12 Education | 11.29 | 1.82 | 11.30 | 1.58 | 646356 |
| Spouse's Postsecondary Education | 1.04 | 1.72 | 0.94 | 1.62 | 646356 |
| Head's Occupational Status | 37.89 | 20.33 | 37.89 | 19.18 | 769035 |
| Spouse's Occupational Status | 39.54 | 18.81 | 39.39 | 13.22 | 420439 |

| Table 3: F-statistics for Nested Model Tests* | | | | |
|---|---|---|---|---|
| Test[a] | Value | df 1[b] | df 2[c] | P-Value |
| Model 1 nested in Model 2 | 3.524 | 1538 | 2864.8521 | 0 |
| Model 1 nested in Model 5 | 2.818 | 2185 | 20734.832 | 0 |
| Model 5 nested in Model 2 | 1.079 | 647 | 223.45052 | 0.251 |

*This output comes from inputting the likelihood ratio statistics into the SAS macro (COMBCHI) written by Allison

[a]Model 1 is estimated with Equation 1, Model 2 is estimated with Equation 2 and so on

[b]This is the numerator degrees of freedom

[c]This is the denominator degrees of freedom

Table 4: Predicted Log Odds of being Retained from the Logit with Partial Interaction Constraints Evaluated at the Grand Means of the Covariates

| Age | Period | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | 1972-1976 | 1977-1980 | 1981-1984 | 1985-1988 | 1989-1992 | 1993-1996 | 1997-2000 | 2001-2003 |
| 6 | -1.5876 | -1.1201 | -0.7788 | -0.3750 | -0.2448 | -0.3827 | -0.3134 | -0.4161 |
| 7 | -0.7568 | -0.5346 | -0.2493 | 0.0234 | 0.1837 | -0.0236 | 0.1705 | -0.0255 |
| 8 | -0.5124 | -0.2990 | -0.0281 | 0.1705 | 0.3035 | 0.0606 | 0.2098 | 0.1599 |
| 9 | -0.4018 | -0.1685 | 0.0556 | 0.2698 | 0.3488 | 0.2593 | 0.2521 | 0.3086 |
| 10 | -0.2482 | -0.1063 | 0.1409 | 0.3225 | 0.4549 | 0.3390 | 0.3456 | 0.3048 |
| 11 | -0.1558 | -0.0368 | 0.1889 | 0.3731 | 0.4809 | 0.4772 | 0.4263 | 0.3963 |
| 12 | -0.1720 | -0.0588 | 0.2000 | 0.3576 | 0.5671 | 0.6039 | 0.4811 | 0.4811 |
| 13 | -0.1116 | -0.1198 | 0.2222 | 0.4581 | 0.5912 | 0.6141 | 0.5693 | 0.4574 |
| 14 | -0.1660 | -0.1027 | 0.0383 | 0.2800 | 0.6208 | 0.6194 | 0.6096 | 0.5089 |
| 15 | 0.0691 | 0.1305 | 0.2238 | 0.3883 | 0.5835 | 0.7125 | 0.6371 | 0.6644 |
| 16 | 0.1970 | 0.3548 | 0.3906 | 0.4791 | 0.6715 | 0.8208 | 0.8997 | 0.7888 |
| 17 | 0.2276 | 0.3639 | 0.3407 | 0.5581 | 0.7396 | 0.8335 | 0.8767 | 0.8225 |

| Table 5: Interaction Constraint Factor and Social Background Characteristic Estimates | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Lambda j | | | Lambda k | | | Beta | | |
| Age | Coef | SE | Period | Coef | SE | Variable | Coef | SE |
| 6 | 0 | -------- | 1972-1976 | 0 | -------- | Male | 0.3163 | 0.0271 |
| 7 | 0.368 | 0.113 | 1977-1980 | -0.040 | 0.066 | Logged Income | -0.1346 | 0.0121 |
| 8 | 0.479 | 0.123 | 1981-1984 | -0.130 | 0.064 | Home Ownership | -0.1791 | 0.0164 |
| 9 | 0.528 | 0.126 | 1985-1988 | -0.132 | 0.064 | Head's K-12 Education | -0.0364 | 0.0034 |
| 10 | 0.619 | 0.132 | 1989-1992 | -0.273 | 0.060 | | | |
| 11 | 0.682 | 0.136 | 1993-1996 | -0.294 | 0.062 | Head's Occupational Status | -0.0026 | 0.0003 |
| 12 | 0.833 | 0.148 | 1997-2000 | -0.559 | 0.069 | | | |
| 13 | 0.693 | 0.137 | 2001-2003 | -0.544 | 0.082 | Spouse's Occupational Status | -0.0021 | 0.0003 |
| 14 | 0.950 | 0.157 | | | | | | |
| 15 | 0.974 | 0.158 | | | | Total Children In Household | 0.0515 | 0.0047 |
| 16 | 1.093 | 0.167 | | | | | | |
| 17 | 1.173 | 0.175 | | | | | | |

| Table 6: Odds Ratios of Being Black versus Non-Hispanic Whites | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Age | Period | | | | | | | |
| | 1972-1976 | 1977-1980 | 1981-1984 | 1985-1988 | 1989-1992 | 1993-1996 | 1997-2000 | 2001-2003 |
| 6 | 0.404 | 0.529 | 0.930 | 0.680 | 0.756 | 0.610 | 0.793 | 0.579 |
| 7 | 0.925 | 0.955 | 1.022 | 0.760 | 0.707 | 0.942 | 0.921 | 0.903 |
| 8 | 0.711 | 0.985 | 0.932 | 1.046 | 0.648 | 0.760 | 0.701 | 0.744 |
| 9 | 1.058 | 0.995 | 1.175 | 0.975 | 1.175 | 0.903 | 0.898 | 0.923 |
| 10 | 1.039 | 1.077 | 1.017 | 1.084 | 0.977 | 0.876 | 1.037 | 0.879 |
| 11 | 1.040 | 1.216 | 1.276 | 1.016 | 1.187 | 0.862 | 1.093 | 0.937 |
| 12 | 0.800 | 0.910 | 1.053 | 1.146 | 1.092 | 0.884 | 1.280 | 1.293 |
| 13 | 0.980 | 0.926 | 1.173 | 1.164 | 1.188 | 1.010 | 1.043 | 0.741 |
| 14 | 1.099 | 1.037 | 1.083 | 1.110 | 1.366 | 1.055 | 1.121 | 1.067 |
| 15 | 0.959 | 0.956 | 0.883 | 1.063 | 0.957 | 0.775 | 1.017 | 1.230 |
| 16 | 0.919 | 0.951 | 1.062 | 1.027 | 1.209 | 1.043 | 1.178 | 1.116 |
| 17 | 1.066 | 1.015 | 1.001 | 1.272 | 1.240 | 1.088 | 0.863 | 0.986 |

**Figure 1: Proportion of Children Below Modal Grade arranged by Birth Cohort**
(Trend lines are 3 year moving averages)

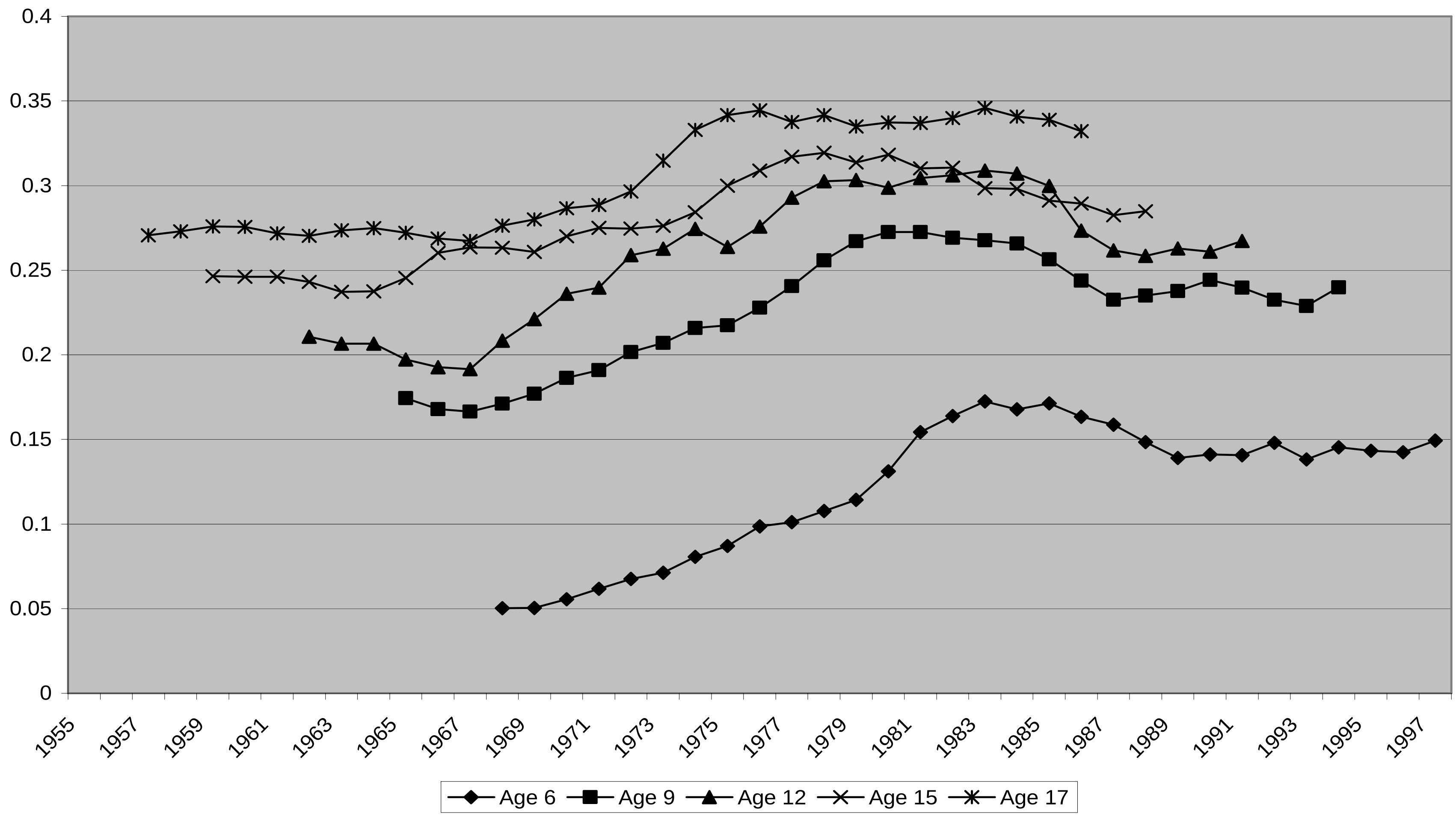


**Figure 2: Proportion of Children Below Modal Grade Adjusting for BMG 6 year olds**
(Trend lines are 3 year moving averages)

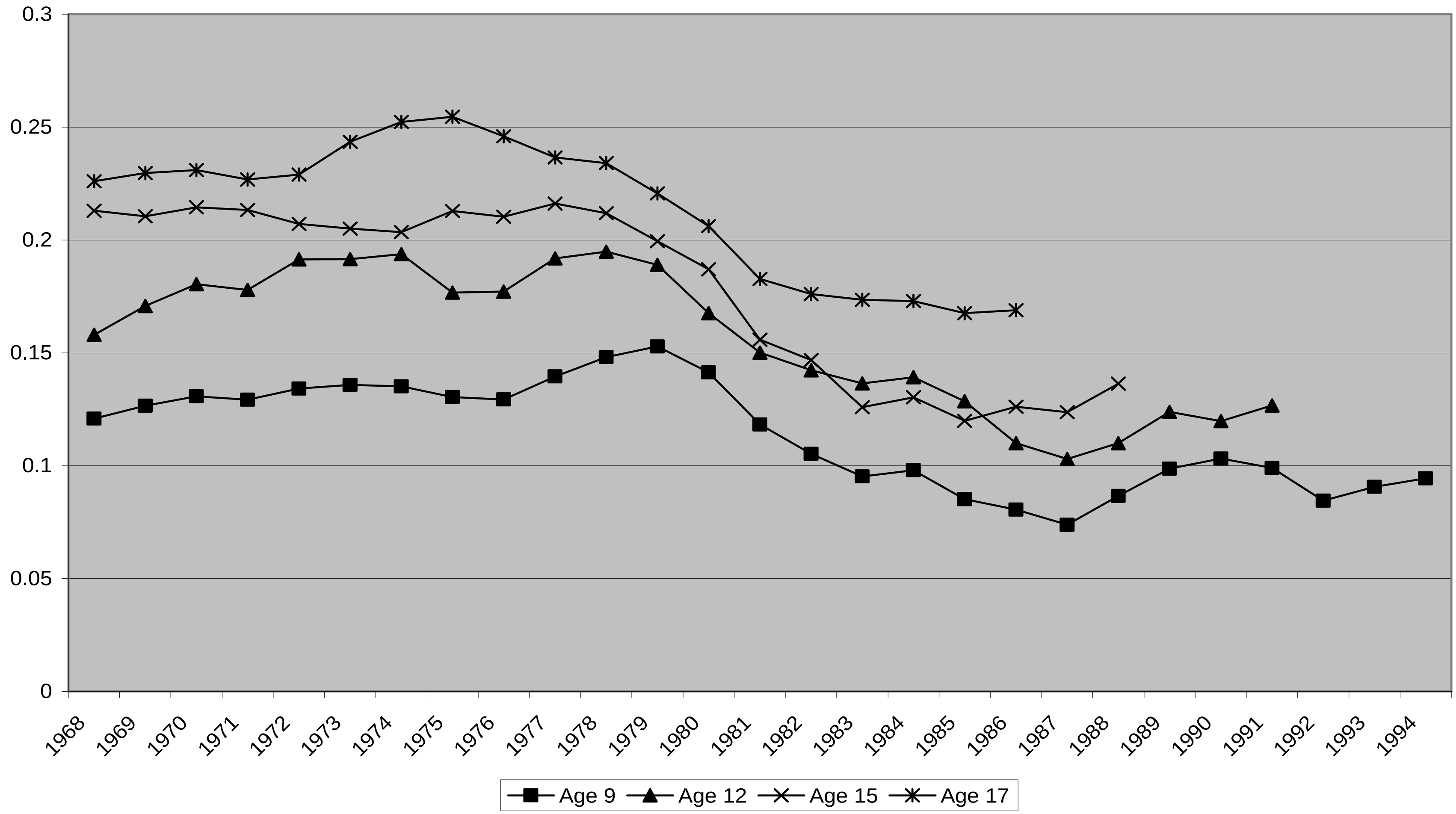
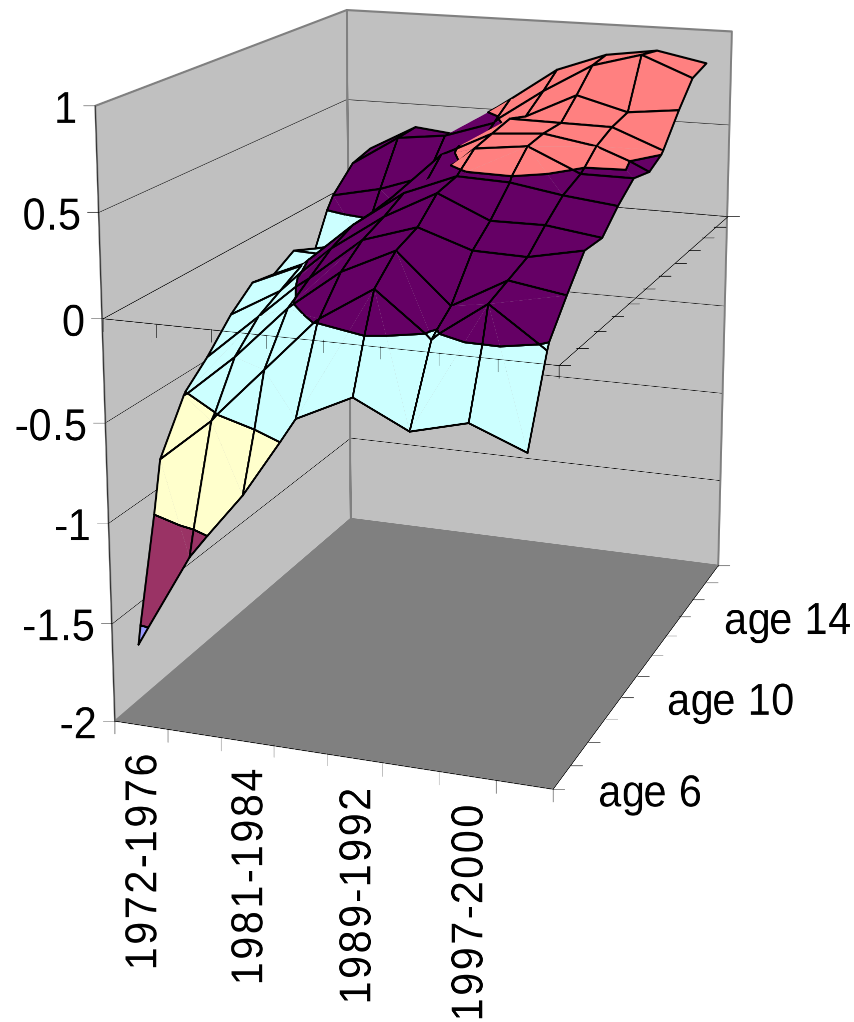


Figure 3: Predicted Log Odds of Being BMG from the LPIC Model



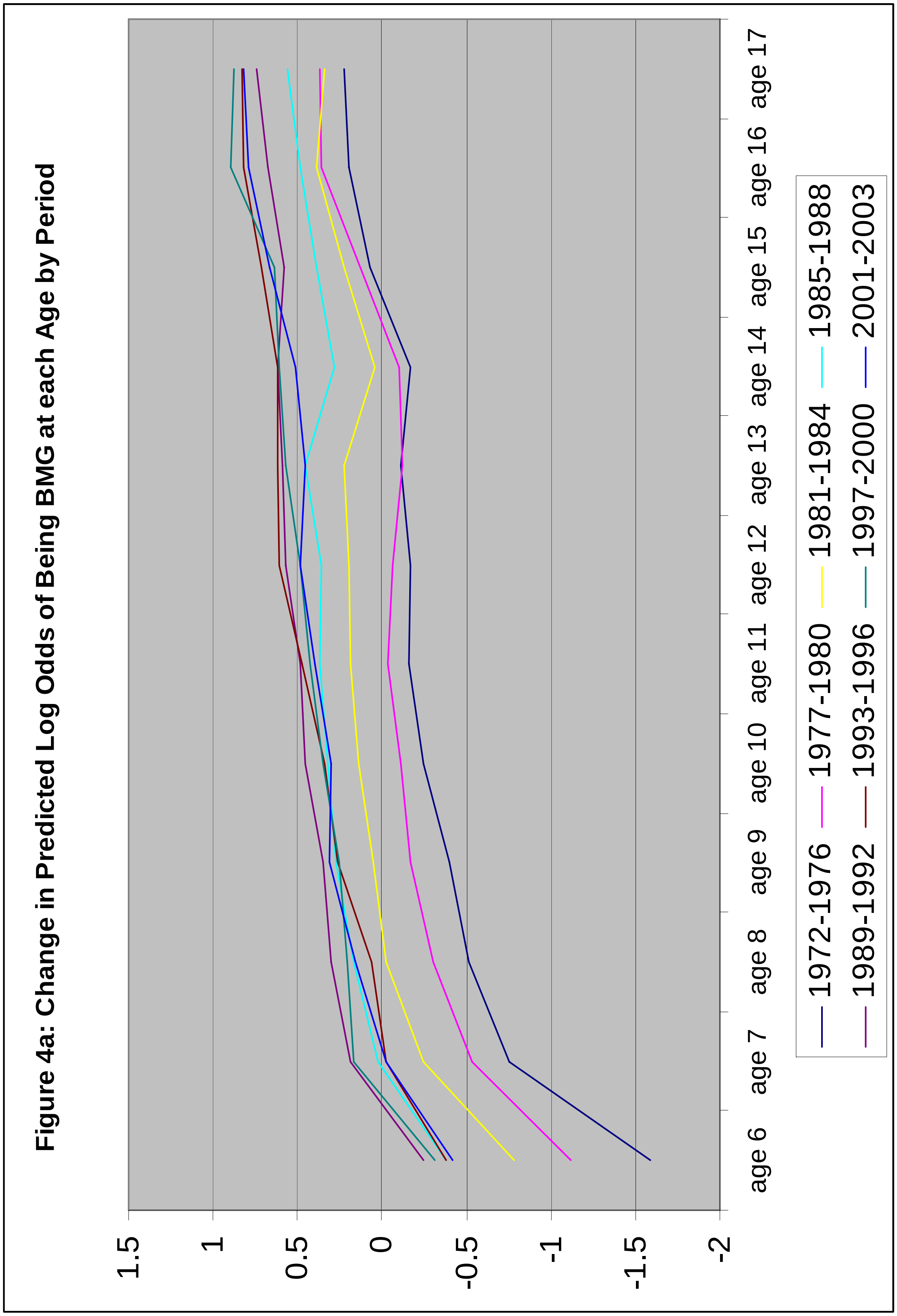
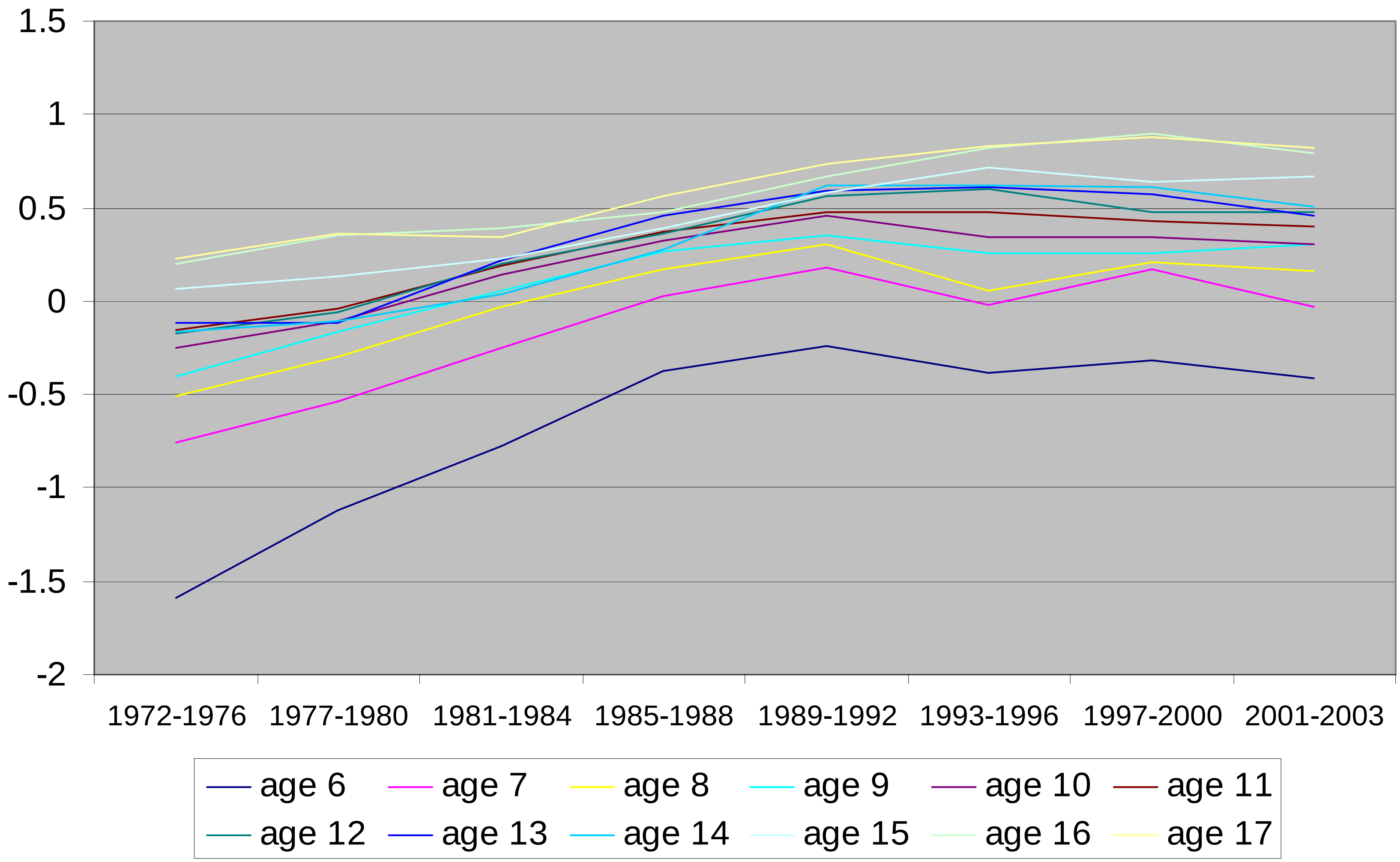



**Figure 4a: Change in Predicted Log Odds of Being BMG at each Age by Period**

**Figure 4b: Change in Predicted Log Odds of Being BMG at each Period by Age**



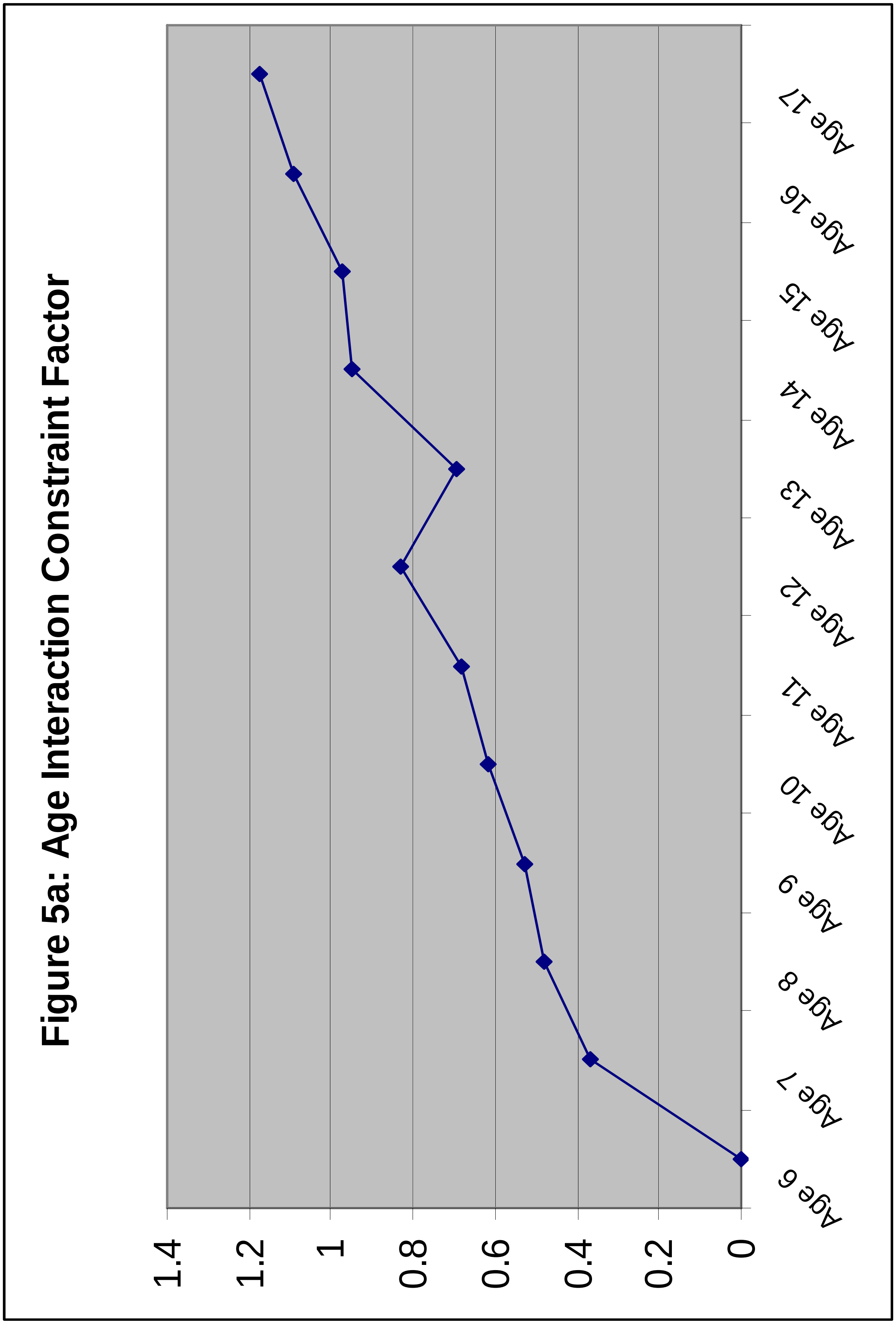


Figure 5a: Age Interaction Constraint Factor

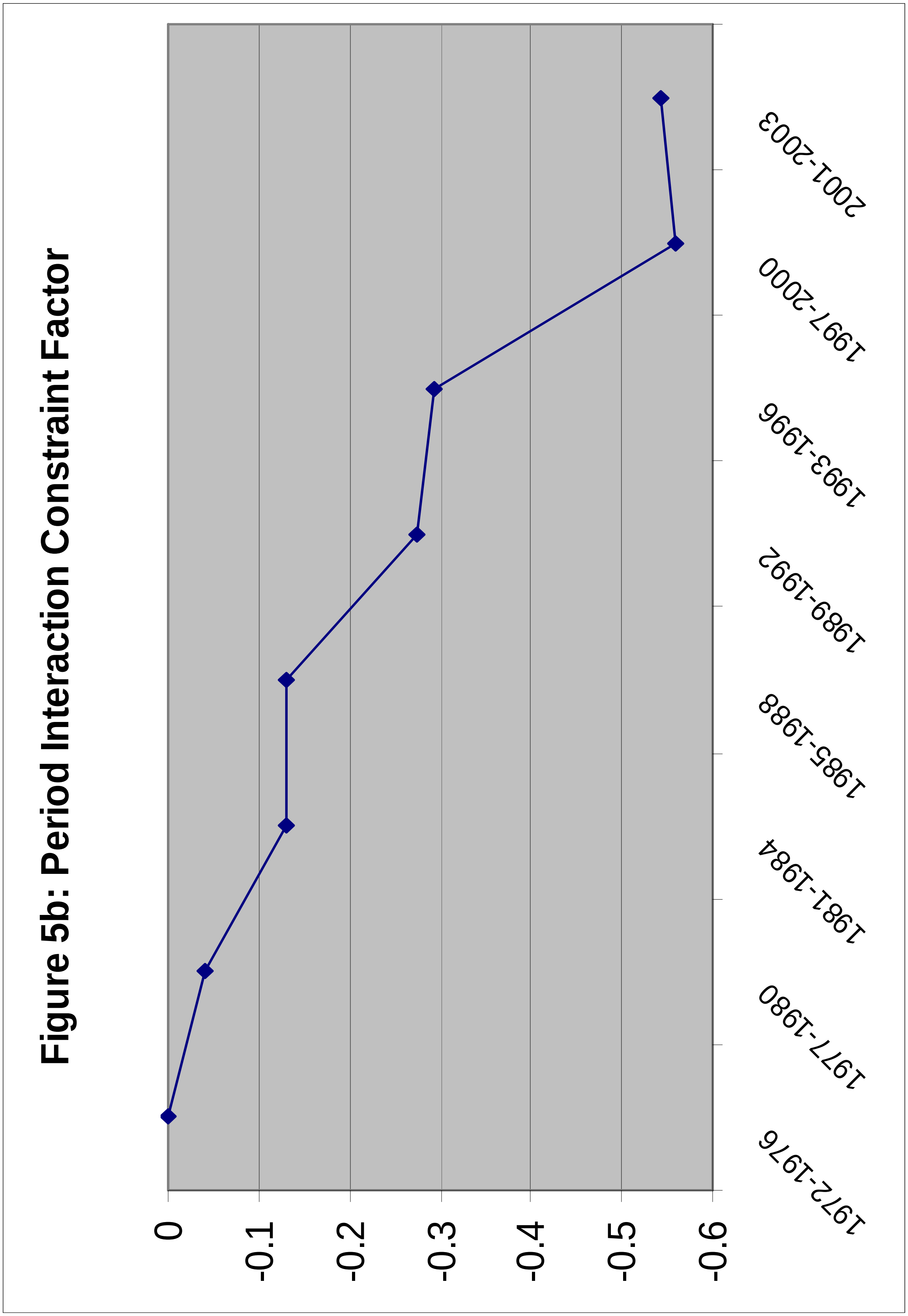


Figure 5b: Period Interaction Constraint Factor
0
-0.1
-0.2
-0.3
-0.4
-0.5
-0.6
1972-1976
1977-1980
1981-1984
1985-1988
1989-1992
1993-1996
1997-2000
2001-2003