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Lifecycle bias in estimates of intergenerational earnings persistence

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Abstract

This paper identifies a significant negative relationship between estimated intergenerational earnings persistence and the age at which fathers are observed. In total, the estimation methodology and the age of the father at observation account for 40 percent of the variation among existing studies. The paper explores two possible causes of this pattern: increasing attenuation bias resulting from growing transitory earnings variance and a lifecycle bias which follows from the rise in permanent earnings variance over the lifecycle. Evidence presented favors the latter explanation over the former. The paper also considers both formal and informal approaches to mitigating the lifecycle bias.

1. Introduction

As intergenerational panel data sets have developed and proliferated, economists have attempted to identify and understand differences in the degree of intergenerational earnings persistence across space and time.[1] For example, Lee and Solon (2004) and Mayer and Lopoo (2004) examine the trend across time; Couch and Dunn (1997),

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[1] Following the convention of the literature, the degree of earnings persistence is defined as the elasticity of son's earnings with respect to father's earnings. Also note that in this paper 'earnings persistence' always refers to *intergenerational* earnings persistence.

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Björklund and Jäntti (1997), and Grawe (2004) explore differences between countries; and Mulligan (1999) studies the effect of differences in school quality across US states. Of course, estimation bias is a widely discussed concern. The mere presence of bias is not always a problem since much of this growing literature is comparative; so long as the bias is equally present in all estimates the comparison remains useful. However, in order to comfortably live with biases it is imperative that we know their sources.

An examination of published estimates of intergenerational earnings persistence suggests the magnitude of the biases could be substantial given the large variation found even among studies based on the same data source and employing similar estimation methods. For example, both Solon (1992) and Couch and Dunn (1997) use Panel Study of Income Dynamics (PSID) data and adjust for errors-in-variables bias, but the estimate from the first study is more than three times that of the second (0.41 compared to 0.13).[2] Clearly some aspect of sample selection is creating significantly disparate results. If we can identify sample selection criteria that *systematically* generate different estimates we will be better able to construct studies that compare the degree of earnings persistence over time, between countries, and across policy regimes. Moreover, we may explain the wide variation found among empirical estimates of intergenerational earnings persistence.

The next section of the paper demonstrates that variation in very basic sample selection criteria explains a substantial fraction of the differences between existing persistence estimates. In particular, there is a strong, negative relationship between the average age of fathers at the point of observation and the earnings persistence estimate. In fact, controlling for father's age at observation explains 20 percent of the variance among studies with similar estimation methodologies.

Drawing on the econometrics literature, Section III examines two possible explanations for this finding. The first explanation, based on Solon (1989, 1992) and others, is that increasing transitory earnings variance over time results in ever-greater errors-in-variables attenuation bias. An alternative explanation is derived from a slight modification of Jenkins (1987), a modification suggested by standard models of the labor market. Human capital accumulation creates greater wage growth for workers with higher lifetime earnings. As a result, deviations of observed earnings from lifetime earnings are correlated with the level of observed earnings. Early in the lifecycle this correlation is negative, but as the cohort ages the correlation ultimately turns positive. The result is a lifecycle bias in earnings persistence estimates that is positive early in the lifecycle, diminishes as fathers age, and ultimately becomes negative-and so earnings persistence estimates diminish with the age of the father. Having identified possible causes, the section then provides several pieces of evidence which point away from the first and toward the second explanation. Section IV discusses possible strategies for dealing with lifecycle bias in earnings persistence estimation and Section V concludes.

[2] The reason the sample selection rules are so different in Couch and Dunn as compared with Solon is that Couch and Dunn match years of observation in the PSID with those available in the German Socioeconomic Panel so that a cross-country comparison can be made.

2. Age-dependence in intergenerational earnings persistence estimates

To determine whether earnings persistence is weak or strong in a particular country, economists often compare estimates for other countries drawn from different studies. (See Solon, 2002; Lefranc and Trannoy, forthcoming; Lillard and Kilburn, 1995 for examples.) As international studies have increased in number, economists have begun to analyze the estimates from these studies in conjunction with theoretical models to uncover the root causes of earnings persistence. For example, Solon (2002) compares resulting estimates from eight countries and employs the human capital accumulation model in Solon (2004) to suggest possible explanations for the differences identified.

Clearly the effectiveness of efforts to document and understand variation in persistence estimates hinges on the comparability of estimates across studies. But the fact that we observe enormous differences between studies from the same data set employing the same estimation methodologies suggests reason for concern. This variation is clear in Table 1 which presents intergenerational earnings persistence estimates from all studies known to the author that meet several requirements. First, all studies in Table 1 mitigate attenuation bias following the suggestions in Solon (1992), either using multi-year averages of earnings as the dependent variable or employing instrumental variables (IV) estimation. Second, the studies impose no extraneous sample selection rules; the samples attempt to measure persistence among all those attached to the labor force in a given population.[3] Finally, papers are excluded from the table if they do not report the mean age of fathers at observation or at least enough information to create a reasonable interval around the mean age.[4] The variation within this group of studies is substantial: standard deviation = 0.13 compared to a mean of 0.27.

Many existing studies have either observed or discussed the potential for a positive association between the age of *sons* at observation and estimates of intergenerational earnings persistence. (See Behrman and Taubman, 1985; Reville, 1995; Couch and Dunn, 1997; Solon, 1999; Chadwick and Solon, 2002, and Solon, 2002. The results in Zimmerman 1992, also exhibit this pattern, though it is not noted by the author.) Probably because these authors focus on the role of transitory earnings variance in biasing persistence estimates (for instance, Björklund (1993) demonstrates, earnings observations are especially noisy prior to age 30), the importance of father's age has been ignored. However, Table 2 shows a strong, negative relationship between the age of father at observation and estimated earnings persistence. Using the data from Table 1, persistence estimates were regressed on a dummy variable noting the manner of

[3] The most notable paper excluded due to this restriction is Zimmerman (1992) which limits the sample to families in which both fathers and sons are employed 30 hours per week, 30 weeks per year. Because earnings persistence increases with father's earnings, this restriction augments Zimmerman's persistence estimate. Altonji and Dunn (1991), included in Table 1, estimate earnings persistence with the same National Longitudinal Study data without the additional hours and weeks restrictions.

[4] When it is particularly difficult to infer the average age of the father, a question mark follows the range. If a paper includes multiple earnings persistence estimates, the estimate included in Table 1 is the one generated when sample selection rules most closely correspond to the selection rules in Solon (1992) – a) positive annual earnings are required in several years which are averaged to control for measurement error and b) only the oldest son available is included.

Table 1
Estimates of intergenerational earnings persistence organized by mean age of father

Author	Mean age of father	Mean year of father observation	Estimate	Location
Lefranc and Trannoy (forthcoming)*,@	34	1964.0	0.41	France
Lillard and Kilburn (1995)	30–40[?]	1975.5	0.27	Malaysia
Björklund and Chadwick (2003)	40.5	1972.5	0.24	Sweden
Corak and Heisz (1999)	40–45	1980.0	0.23	Canada
Mulligan (1997)	40–45	1969.0	0.33	US
Björklund and Jäntti (1997)*,#	43	1970.2	0.28	Sweden
Shea (2000)**	44	1969.0	0.36	US
Solon (1992)	44	1969.0	0.41	US
Björklund and Jäntti (1997)*,#	45	1969.0	0.42	US
Mazumder (forthcoming)[a]	46	1982.0	0.39	US
Peters (1992)	47	1969.5	0.14	US
Bratberg et al. (forthcoming)[b]	47	1978.0	0.12	Norway
Dearden et al. (1997)*	45–50	1974.0	0.58	UK
Tsai (1983)	45–50[?]	1958.5	0.28	Wisconsin
Österbacka (2001)	48.5	1972.5	0.13	Finland
Couch and Dunn (1997) #	51	1986.5	0.11	Germany
Wiegand (1997)*	51	1984.0	0.20	Germany
Altonji and Dunn (1991)	52	1967.3	0.18	US
Couch and Dunn (1997) #	53	1986.5	0.13	US
Österberg (2000)	53	1979.0	0.13	Sweden

[a] Mazumder's estimate using three years of father's data is chosen in order to most closely match the methodology of the other studies in the table. When he uses six years of father's earnings data, his estimate is 0.47. His estimates resulting from ten and fifteen years of data are avoided since these estimates were found to be very sensitive to the treatment of top-coded earnings reports.

[b] The 1960 cohort is chosen from the Bratberg et al. study since it is twice as large as the 1950 cohort sample. The regression fit in Table 2 would be tighter if the 1950 estimates were used.

* Studies using IV estimation.

** Because Shea does not include the same number of earnings observations for each father, the precise years of observation and the average year of observation are clear from the study. However, the data are intended to be very similar to that of Solon.

[?] The range attributed to Mean Age of Father is particularly difficult to infer from information in the paper.

[#] Studies which are expressly cross-country comparisons.

[@] The estimate found is that when sons are measured in 1993 and fathers in 1964. Average age of fathers taken from personal correspondence with authors.

attenuation bias correction (=1 if employing IV, =0 if using a multi-year average of father's earnings). The results are reported in column 1 of Table 2. Then the age of father at observation was added to the regression; column 2 documents a substantial negative effect of father's age. Observing fathers at age 53 as opposed to age 34 (the range of observation among studies in Table 1) reduces earnings persistence estimates by 0.18 (*p-value* = *0.062*).[5] After controlling for the method of estimation, the age of

[5] To the extent that I have erred in approximating the age at which father's earnings is observed in the five cases where an exact report was not available, I have introduced measurement error into the analysis which presumably serves only to reduce the explanatory power of this variable.

Table 2
The effect of father's age on estimates of earnings persistence found in 20 existing studies

Dependent variable: estimated earnings persistence	(1)	(2)	(3)
IV dummy (1 if study uses IV to control for attenuation bias; 0 otherwise)	0.148 (2.477)	0.128 (2.274)	0.136 (2.380)
Father's age		− 0.009 (1.997)	
Year of father observation			− 0.006 (1.744)
R-square	0.254	0.396	0.367

Regression (1) examines variation in estimated earnings persistence by method of correction for attenuation bias (instrumental variables correction vs. averaging father's earnings over multiple years). Regression (2) adds father's age to the analysis. Regression (3) replaces father's age with the year of father observation to explore whether age-effects may actually be year- effects. Absolute t-statistics in parenthesis.

father at observation accounts for 20 percent of the variation among studies.[6] (The method of error correction and the mean age of fathers combine to explain 40 percent of the total variation.)

In addition to explaining much of the variation between studies, these results substantially alter perceptions of 'outliers'. For instance, without considering the age of fathers, the estimates of Couch and Dunn (1997) for Germany and Österbacka (2001) for Finland (both around 0.1) appear to be extraordinarily low. However, considering the fact that both studies observe fathers late in the lifecycle, the results appear in line with other studies.

One alternative explanation for the observed age-dependence is that the age variable is picking up effects properly attributed to the year of observation because, in the studies of Table 1, the age of father at observation is positively correlated with year of observation ($D = 0.41$). If earnings persistence is decreasing with time, this could look like negative age-dependence. Column 3 of Table 2 reports regression results replacing the age of father with the year of observation to test this hypothesis. Estimated earnings persistence is negatively related with the year of father observation, but the relationship is less significant in both practical and statistical terms (*p-value* = *0.099*). Thus, we cannot rule out the hypothesis that what appear to be age effects are time effects instead. However, it should be noted that for the US, direct examinations of changes in persistence over time have not found a clear trend. Levine and Mazumder (2002) (PSID sample), Fertig (forthcoming), and Mayer and Lopoo (2004) find a statistically insignificant decline in earnings persistence over time; on the other hand, Levine and Mazumder (2002) (National Longitudinal Study sample) find a statistically significant *increase* over time. And Lee and Solon (2004) report no change. While it appears reasonable to interpret the results of Table 2 as what they appear to be–a negative relationship between the age of father and estimated earnings persistence-a closer examination is warranted beginning with a better understanding of the possible sources of age-dependence in persistence estimates.

[6] IV estimates are higher by 0.13 on average suggesting either that multi-year measures of father's earnings fail to entirely eliminate measurement error and/or that the instruments used are endogenous. This is consistent with findings in Solon (1992) and many other studies.

3. Possible explanations for age-dependence in earnings persistence estimates

To examine the source of the observed age-dependence in more depth, this section presents a simple model of intergenerational earnings transmission based on Jenkins (1987). The model provides the basis for two explanations for the age-dependence found in the previous section. Both explanations rest on biases in estimates of intergenerational earnings persistence. After presenting the model and examining the sources of bias, evidence for and against the alternative explanations is considered.

3.1. Attenuation bias and lifecyle bias

Jenkins (1987) presents a simple model in which both parent and child work for two periods. The log earnings of fathers are

$$
\begin{aligned}
F_1 &= \alpha_{F1} G_F + v_1 \\
F_2 &= \alpha_{F2} G_F + (1+\delta) v_1 + v_2 v_1 \perp G_F, v_2 \perp G_F, v_1 \perp v_2, \\
E_{(v_1)} &= E(v_2) = 0, \alpha_{F1}, \alpha_{F2} > 0, \frac{(\alpha_{F1} + \alpha_{F2})}{2} = 1 \qquad (1)
\end{aligned}
$$

where F_i is log earnings in period i. Log earnings in a single period can be broken into permanent and transitory components. First, $\alpha_{Fi} G_F$ for $i=1,2$ represents the annual permanent component of earnings in year i. Theory and empirical work suggest that earnings growth is positively correlated with the level of earnings. In the model, this is captured by the assumption that α_F increases with age.[7] Second, v_1 represents the transitory earnings component in year 1 while $(1+\delta)v_1+v_2$ is transitory earnings in year 2. $\delta>0$ allows for persistence in transitory earnings. Because the transitory components of log earnings have an expected value of zero, the expected total lifetime log earnings $E(F_1+F_2) \equiv E(F)=G_F$. For this reason G_F will be called lifetime permanent earnings.

The log earnings of sons follow an analogous system, though the rate at which the variance of annual permanent log earnings increases may differ by generation (that is, $\alpha_{S1} \neq \alpha_{F1}$ and $\alpha_{S2} \neq \alpha_{F2}$). Given the increase in returns to education experienced in many countries in recent decades, we might reasonably assume that the rate of growth in annual permanent log earnings variance is greater for sons than fathers. That is, $\alpha_{S2} - \alpha_{S1} > \alpha_{F2} - \alpha_{F1}$. Fig. 1 depicts functions α_S (age) and α_F (age) capturing this feature in the more realistic case of many-period lifetimes for both father and son.

While the transitory components of log earnings are assumed independent across generations, lifetime permanent earnings is connected according to

$$G_S = g G_F + e \qquad (2)$$

where e is a random variable which is independent of all other variables. Depending on their purpose, many economists wish to estimate one of two measures of intergenerational earnings association. The first measure of interest is the structural intergenerational association of lifetime permanent earnings g. The second sought after parameter is the

[7] Jenkins assumes the reverse in his simulations as discussed in the next section. If we wish to also allow earnings to grow, on average, with age then a constant term could be added to the expression for F_2. Because this addition will not alter any of the discussion in this paper it is omitted.

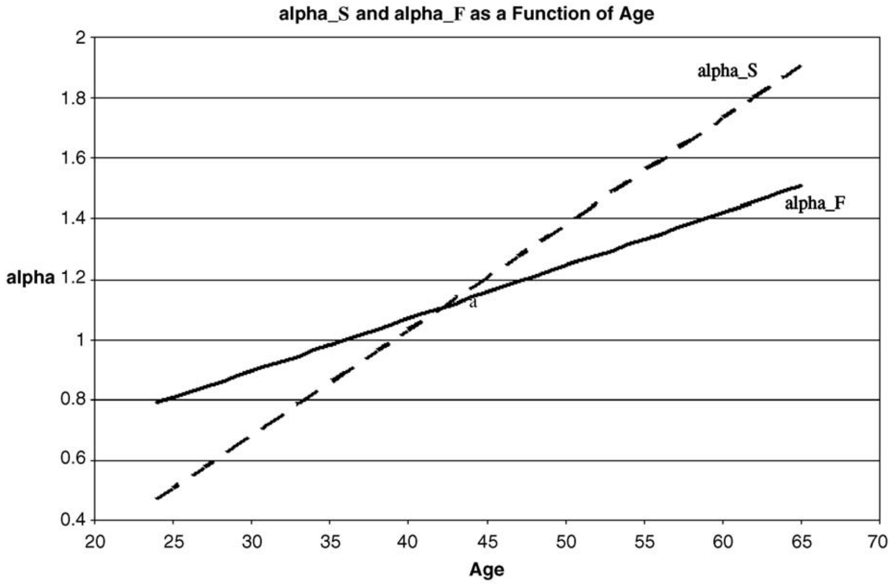

Fig. 1. How annual permanent log earnings variance parameters change over the lifecycle in two generations.

empirical relationship between the total lifetime log earnings of father and son-the regression coefficient when father's total lifetime log earnings (F) are regressed on son's total lifetime log earnings (S):

$$\beta = \frac{Cov(S,F)}{Var(F)} = g\frac{(\alpha_{S1}+\alpha_{S2})(\alpha_{F1}+\alpha_{F2})\sigma_G^2}{(\alpha_{F1}+\alpha_{F2})^2\sigma_G^2+(1+\delta)^2\sigma_{v1}^2+\sigma_{v2}^2} \tag{3}$$

where $\sigma_G^2 = var(G_F)$ and $\sigma_{\upsilon i}^2$ represents the variance of transitory earnings shocks among fathers in period i. If there were no transitory earnings shocks, then $\beta = g$ because $\alpha_{S2}+\alpha_{S1}=\alpha_{F2}+\alpha_{F1}$. However, with transitory earnings present the elasticity of total lifetime earnings (the elasticity of S with respect to F or β) falls short of the elasticity of lifetime permanent earnings (the elasticity of G_S with respect to G_F or g). This is easily represented as a classical errors-in-variables attenuation bias if total lifetime log earnings (F and S) are viewed as error-ridden measures of lifetime permanent earnings (G_F and G_S).

Economists must estimate g or β based on limited data-much less than a lifetime of data for both fathers and sons. In Jenkins' model this is the equivalent of observing log earnings for each generation in only one of the two periods (either F_1 or F_2 for fathers and either S_1 or S_2 for sons). Jenkins shows how ordinary least squares (OLS) estimates based on these limited 'snapshots' are biased no matter which combination of periods of father and son are observed. Consider the case where sons and fathers are observed only in periods i and j respectively, where $i,j \in 0\{1,2\}$. The probability limit of the OLS estimate based on single-period, snapshot measures of earnings is

$$\hat{\beta} = g\frac{\alpha_{Si}\alpha_{Fj}\sigma_G^2}{\alpha_{Fj}^2\sigma_G^2+\sigma_{\upsilon j}^2+\delta^2\sigma_{\upsilon 1}^2(j-1)} \tag{4}$$

which is particular, is neither g or β.

Two biases explain the deviation of from $\hat{\beta}$ and β. The first pertains to the role of transitory earnings. When g is the parameter of interest, this bias is most easily understood as a classical errors-in-variables attenuation bias. As noted in Atkinson (1980-81) and Atkinson et al. (1983) and detailed in Solon (1989, 1992) and Mazumder (forthcoming) among others, classical measurement error in father's earnings produces an attenuation bias that reduces estimated earnings persistence. In Eq. (4) this is seen in the presence of $\sigma_{\upsilon j}^{2}+\delta^{2}\sigma_{\upsilon 1}^{2}(j-1)$ in the denominator. This problem is well-noted in the literature with substantial empirical work demonstrating its relevance (see Solon, 1992; Zimmerman, 1992; Couch and Dunn, 1997; Mazumder, forthcoming for examples). IV estimation eliminates this bias.

When β is the parameter of interest, the bias results not from too much transitory earnings variance, but too little. In any given period, only one period's transitory earnings are present. In other words, one of the terms in the sum $\sigma_{\upsilon j}^{2}+\delta^{2}\sigma_{\upsilon 1}^{2}(j-1)$ is missing. In this case IV estimation, far from eliminating bias, contributes its own bias by eliminating transitory earnings variance from the probability limit altogether. (As β is the parameter of interest in Jenkins 1987, this problem is noted in that paper.)

A second source of bias, a lifecycle bias which results from the fact that the annual permanent component of earnings ($\alpha_{Fi}G_F$) does not generally equal lifetime permanent earnings (G_F), is more difficult to correct.[8] This presents similar problems for the estimation of both g and β. Even when transitory earnings are non-existent (or when IV estimation is employed to address attenuation bias) the OLS (or IV) probability limit equals $g^*(\alpha_{Si}/\alpha_{Fj})$.[9] Because α_{Si} generally does not equal α_{Fj}, estimates of intergenerational earnings persistence based on a short span of earnings data ('snapshots' in Jenkins' terminology) are generally biased.

Jenkins performs simulations to evaluate the degree to which attenuation and lifecycle biases combine to bias persistence estimates based on earnings 'snapshots' using β as a benchmark. He concludes that the total bias may be large depending on the parameter values and that "no obvious *general* rules about bias...can be made" to ensure unbiased estimates (p. 1152, emphasis in original).[10] The remainder of this section demonstrates that even though lifecycle and attenuation biases substantially affect estimates we may yet predict how the biases change across the lifecycle. So even if no general rule leads to unbiased estimates, general rules can be constructed allowing us to compare more meaningfully estimates of earnings persistence across studies.

3.2. Connections between age-dependence and attenuation and lifecycle biases

The attenuation and lifecycle biases embodied in Eq. (4) can be used to produce two possible explanations which individually or combined account for the age-dependence in

[8] The term 'lifecycle bias' has been used widely in reference to a variety of different things. In particular, Jenkins uses the term to refer to the combined effect of attenuation bias and what is termed a 'lifecycle bias' in this paper.

[9] The lifecycle bias might also be thought of in terms of a measurement error problem: Expected earnings in a given period ($\alpha_i G_i$) differ from lifetime permanent earnings (G_i). The reason why IV fails to address the resulting bias is that in this case the measurement error ($\alpha_i G_i - G_i$) is correlated with the dependent variable (G_i). See Kane et al. (1999) and Bound and Solon (1999) for discussion of this econometric issue in other contexts.

[10] Jenkins also considers the effects of within-generation age differences under the assumption that earnings grow/diminish with age at a rate that is constant across individuals.

earnings persistence estimates found in Section II. The explanation based on attenuation bias presumes that transitory earnings variance has increased over time producing ever larger downward bias and so ever lower earnings persistence estimates. (An alternative hypothesis is that transitory earnings variance increases over the lifecycle. However, most economists believe that the transitory component of earnings actually decreases in importance over the lifecycle. See Björklund (1993) or Baker and Solon (2003), for example.) Solon (1992) employs a similar argument to explain why earnings persistence estimates are lower when fathers are observed in 1970-71 than when fathers are observed in 1967-69. In that case, the author points to the possible change in transitory earnings variance over the business cycle. Here we consider a longer secular trend of increasing transitory earnings variance, but the argument is the same.

It is also possible to explain the observed age-dependence in terms of a lifecycle bias. To make this connection, we join the work of Jenkins with standard models of human capital accumulation. It is a well-known result of models like Ben-Porath (1967) that earnings growth is positively associated with the level of lifetime earnings. As a result, over most of the lifecycle, the variance of the annual permanent component of earnings ($\alpha_F G_F$) increases with age. In the notation of Jenkins' model, this implies $\alpha_{j1} < \alpha_{j2}$ for $j = S,F$. Jenkins does not make this connection, making no note that α varies *systematically* across age. Moreover, in his simulations he only considers cases in which $\alpha_1 > \alpha_2$ in conflict with both the theoretical and empirical literature. (See Tables 1 and 2 in Jenkins.) This may be one reason why an important result in that paper-that changes in the variance of annual permanent earnings (that is, var($\alpha_F G_F$)) matter as much to earnings persistence estimates as changes in transitory earnings variance-has been underappreciated in the subsequent literature.

Having made the connection with the wider literature, it is now possible to move beyond the limiting conclusion that lifecycle bias matters to a more constructive conclusion: lifecycle bias varies predictably across age. Extend Jenkins' model to allow for many life periods for father and son: at time t_F the permanent component among fathers is given by $\alpha_F(t_F)G_F$ where $\alpha_F(t_F)$ increases with t_F as depicted in Fig. 1. An analogous (though not identical) relationship holds for sons. Suppose we happen to observe sons at the age at which $\alpha_S(t_S) = 1$. (It is easy to work out the pattern of lifecycle bias for other son observation ages.) If fathers are observed early in the lifecycle, the lifecycle bias is large and positive. As the age of father observation increases, the lifecycle bias diminishes. At some point at midlife $\alpha_F(t_F) = \alpha_S(t_S) = 1$ and lifecycle bias is zero. As the observation age for fathers continues to increase, the lifecycle bias becomes increasingly negative. By contrast, when sons are young the bias is negative and then it increases as sons age. As a result, models of human capital accumulation such as Ben-Porath (1967) predict a negative (positive) relationship between the age of father (son) at observation and estimated earnings persistence.

3.3. Evaluating attenuation bias and lifecycle bias as sources of age-dependence

Having shown that it is possible that either attenuation or lifecycle bias explains the agedependence of earnings persistence found in Section II, next consider facts that may help discern which factor is of greater importance. Evidence found undermines the attenuation bias and supports the lifecycle bias as the explanation.

First, consider the hypothesis that attenuation bias has increased over time due to an increase in transitory earnings variance. As noted previously, this hypothesis is directly refuted by the literature searching for time trends in intergenerational earnings persistence. (See Fertig, forthcoming; Lee and Solon, 2004; Levine and Mazumder (2002); Mayer and Lopoo, 2004) If the observed effect of age were actually reflecting an increase in attenuation bias over time, one would expect such a trend to show itself in these studies.

More damaging evidence, however, is found when one considers the premise of the argument. If attenuation bias has increased with time, then it must be that transitory earnings variance has increased over time *relative to* variance in non-transitory earnings. While it is true that transitory earnings variance has increased in North America, permanent earnings variance has actually increased at an equal or faster rate.[11] In Canada, Baker and Solon (2003) find that transitory earnings variance has increased at a somewhat slower rate than permanent earnings variance. In the US, Gottschalk and Moffitt (1994) find that transitory earnings variance grew at a rate between 2/3's and equal to that of permanent earnings variance while Haider (2001) finds equal growth. Baker and Solon (2003) further argue that the US estimates likely overstate the relative increase in transitory earnings variance due to model specification error.) If growing attenuation bias were the cause of the observed age-dependency, then transitory earnings variance would have to have grown at a faster rate than permanent earnings variance and the data contradict this.

The evidence is more favorable toward the lifecycle bias explanation. Of course, it is well known that the premise of the argument is supported by the data-permanent earnings variance does grow with age as Mincer (1958) predicted. Moreover, as is shown next, a study of age-dependency *within* countries produces results consistent with the lifecycle bias interpretation. Specifically, if lifecycle bias contributes significantly to age-dependence in earnings persistence estimates across studies, then we should expect to find two facts. First, age-dependence should be observed in many countries because lifecycle bias results from a fundamental feature of the human capital accumulation process. In particular, it should be seen in countries which experienced dramatic, modest, and limited increases in transitory earnings variance (like the US, Canada, and Germany respectively). Second, when estimates of intergenerational earnings persistence are compared with the age of sons at observation, the relationship should be positive-the reverse of the pattern observed among fathers.

To my knowledge, only Reville (1995) performs a detailed study of age-dependence and then only for sons and only in the American PSID. Zimmerman's (1992) National Longitudinal Survey (NLS) results can be used to identify age-dependence in both fathers and sons, however the fact that Zimmerman restricts his analysis to full-time labor force participants may affect his findings. The limited years covered in the NLS also limit its use as a definitive source on agedependence in earnings persistence estimates. In this paper, age-dependence in both fathers and sons is examined using data from the American PSID (1968–1993) and NLS (1966–1981), the Canadian Intergenerational Income Data (IID) (1978–1998), and German Socioeconomic Panel (GSOEP) (1984–2001). See Appendix A for detailed sample descriptions.

[11] It should also be noted that the attenuation bias explanation is North American-centric The recent rise in transitory earnings variance has not been experienced uniformly throughout the world. Yet, the age-dependence exhibited in Table 2 includes studies in Europe and Asia.

Each data set contains multiple observations for both sons and fathers.[12] Each earnings observation for the son is regressed in turn on each of the available earnings observations for the father including controls for both age and age-squared for both the father and the son.[13] Note, however, that the inclusion of age controls only corrects for changes in *mean* earnings across age. A lifecycle bias remains so long as there are changes in earnings *variance* over the lifecycle.

Eq. (5) presents an example regression in which son's log earnings measured in 1993 are regressed on father's log earnings measured in 1987.

$$y_{s,93} = \lambda + \gamma_1 \text{age}_{s,93} + \gamma_2 \text{age}_{s,93}^2 + \gamma_3 \text{age}_{f,87} + \gamma_4 \text{age}_{f,87}^2 + b y_{f,87} + \varepsilon \qquad (5)$$

In this regression b measures persistence in earnings across generations. In each data set except the GSOEP, each year of father (son) observation can be matched with multiple years of son (father) observation. Earnings persistence estimates are computed for each possible father-son observation pair as described above. For instance, there are five NLS son observations that can be paired with each father observation. So, for each year of father observation, there are five available estimates of earnings persistence. These multiple earnings persistence estimates are then averaged. This average estimated earnings persistence for the particular year of father (son) observation is recorded in Fig. 2 (3) which plots how the average of earnings persistence estimates varies as the age (and year) of father (son) observation increases.[14]

Fig. 2 shows that within each data set the average estimated earnings persistence drops noticeably as the age of father at observation increases. The magnitude of the change is similar in all countries-a little more than one percentage point per year. Excluding the NLS results, Fig. 2 appears as consistent with year effects as with age effects. However, as noted above the broader evidence questions the year-effects interpretation. The fact that all three countries and all four samples exhibit a negative relationship between estimated earnings persistence and father age despite radically different transitory earnings variance evolutions across countries is consistent with the lifecycle bias. Thus, the first prediction of the lifecycle bias is confirmed in the data. (Theory does not predict that the degree of age-dependence across studies should be the same as within all countries. But it seems reasonable to assume that α_F grows at roughly similar rates across countries which means that the degree of age-dependency should be roughly similar across countries. Thus the

[12] In order to avoid confounding effects of sample attrition due to the retirement of fathers, sample selection is limited by the age of fathers. For instance, in the examination of the PSID, fathers are no older than 46 in 1967 so that they are no older than 60 in 1981, the final year of observation. In other words, the graphs below follow a cohort of families across time.

[13] Obviously, single-year measures of earnings contain measurement error and so the levels of earnings persistence estimated in this section are lower than the true degree of persistence. However, in identifying the importance of a life cycle bias, we are interested in the *trend* in estimates over the lifecycle – not the level of persistence itself. This trend is easier to identify when we have a large number of estimates from a wide range of ages. When the analysis is repeated using three-year averages of earnings, the same qualitative results obtain. But with one-third the number of independent persistence estimates, it is more difficult to determine whether the pattern constitutes a trend.

[14] The individual persistence estimates are available from the author on request. The standard errors for the individual estimates are 0.05–0.10 in the NLS and PSID, 0.10–0.19 in the GSOEP, and 0.006–0.009 in the IID.

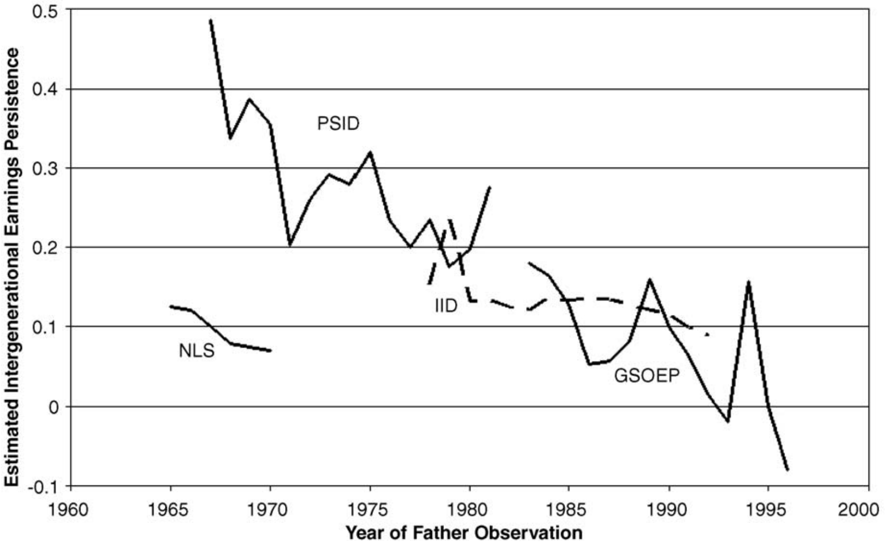

Fig. 2. Variation in earnings persistence estimates as fathers age in the PSID, NLS, IID, and GSOEP data sets.

fact that estimated degree of earnings persistence decreases by approximately one percent for each year of father age both across studies (Table 2) and within the US (PSID and NLS), Germany, and Canada is also consistent with the lifecycle explanation.)

Fig. 3 turns the focus to changes in persistence estimates across the age of son at observation. The positive relationship found by Reville (1995) is again found in the PSID

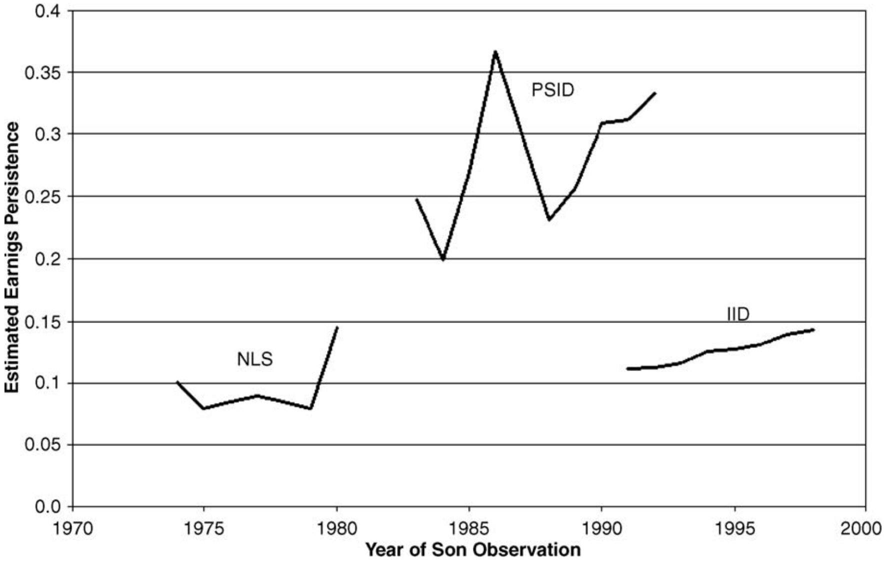

Fig. 3. Variation in earnings persistence estimates as sons age in the PSID, NLS, and IID data sets.

and in the NLS and Canadian IID as well. This confirms the second prediction stemming from the lifecycle bias.

In total, the evidence appears to support the hypothesis that lifecycle bias and not growing attenuation bias is the cause of the negative relationship between age of father and estimated earnings persistence found in Table 1. This is not to say that attenuation bias is unimportant in these studies; undoubtedly all of the non-IV estimates suffer from this downward bias. While Jenkins may be correct that persistence estimates based on single-year 'snapshots' of earnings are inevitably affected by bias, thinking about attenuation and lifecycle bias separately nevertheless leads to useful insights. Even though we may never know the precise magnitude of either bias, well studied facts concerning changes in the variance of transitory and permanent earnings make it easy to predict how the biases will vary with age of father.

4. Correcting for lifecycle bias

As Jenkins notes, estimates of β necessarily require a full lifetime of data because it is otherwise impossible to estimate the degree of transitory earnings variance across the lifecycle.[15] Given the wide interest in estimating g demonstrated by the many works that employ measurement-error corrections, it is reasonable to wonder whether we might yet be able to correct for the bias when estimating this structural parameter. The model of the previous section suggests several related approaches, both formal and informal. Ultimately, formal correction requires strong assumptions coupled with an estimation process that magnifies standard errors. Thus, rough rules of thumb may serve as useful as formal corrections in substantially reducing bias.

Looking back at the previous section, it is clear that single-year measures of earnings produce a systematic bias in persistence estimates, a bias that is positively correlated with father's age and negatively correlated with son's age. Unlike the problem of classical errors-in-variables, the measurement issue pertaining to the lifecycle bias is as relevant to son's earnings (the dependent variable) as it is to father's earnings (the independent variable). Traditional approaches like IV fail; in particular, the IV probability limit is $g^*(\alpha_{Si}/\alpha_{Fj})$. If we knew precisely how earnings variance evolved over the lifecycle for both father and son we could choose a combination of father and son ages such that the bias introduced by the mis-measurement of one was exactly offset by the bias introduced by the mis-measurement of the other. Assuming classical errors-in-variables attenuation bias is corrected for using IV, we seek ages for father and son where $\alpha_S(t_S)=F(t_F)$. Alternatively, as Haider and Solon (2004) point out, knowing $\alpha_S(t_S)$ and $\alpha F(t_F)$ for any particular pair of observation dates we may multiply the earnings persistence estimate by $\alpha_S(t_S)/\alpha_F(t_F)$ to produce a corrected estimate.

Both $\alpha_S(t_S)$ and $\alpha_F(t_F)$ can be estimated if we had data on lifetime permanent earnings and single-year earnings for a sample of sons and a sample of fathers. Recall,

[15] Of course, one could assume that the degree of transitory earnings variance during the unobserved periods were the same as that in the observed periods, but that would simply be to assume the answer.

in the simple case with no transitory earnings that annual log earnings for person i at time t_i equal $\alpha_i(t_i)G_i$ for $i=S,F$ where G_i represents lifetime permanent earnings. Regressing single-year earnings on lifetime earnings for sons yields an estimate for $\alpha_S(t_S)$. If measurement error were not a concern, the same strategy applied to data for fathers would yield an estimate for $\alpha_F(t_F)$. Haider and Solon (2004) extend the analysis to the more likely case in which measurement error is present showing that a "reverse regression" of lifetime earnings on the single-year earnings observation for fathers yields something like $1/\alpha_F(t_F)$ except that it accounts for attenuation bias in addition to lifecycle bias in father's earnings.

Of course, this discussion entirely hypothetical. For if we had enough data to actually observe lifetime permanent earnings G_i we wouldn't have to worry about lifecycle bias in the first place; we would simply use the lifetime earnings in our analysis. In practice, a much shorter panel of earnings combined with a parametric assumption for the earnings trajectory allow us to estimate the shape of earnings over the lifecycle. These estimates in turn can be used to estimate lifetime earnings. Unfortunately, in some data sets only one observation of earnings is available. (This is true of the fathers in the British National Child Development Survey (NCDS), for instance.) In such cases, the estimated earnings trajectory from another data set must be used and the researcher must additionally assume that the age-earnings profile is the same in the two data sets. Next, the estimates of permanent lifetime earnings are used to estimate $\alpha_S(t_S)$ and $\alpha_F(t_F)$. Finally, these estimates can be used to correct the initial earnings persistence estimate in the third estimation stage.

Clearly this approach will significantly magnify standard errors of earnings persistence estimates because the final estimate is the result of many intermediate parameter estimates.[16] (See Murphy and Topel 1985, for a discussion of multi-step estimators.) Given the resulting standard errors, researchers with modest sample sizes may find that such formal corrections are not significantly better than more modest rules of thumb reflecting what we know of lifecycle bias. Fig. 1 suggests two possible strategies for mitigating (though not entirely eliminating) lifecycle bias. Because $\alpha_i(t_i)$ for $i=S,F$ is an increasing function, initially less than one and eventually greater than one, it is reasonable to attempt to measure both fathers and sons near midlife when $\alpha_S(t_S) \cong \alpha_F(t_F) \cong 1$. Combined with standard approaches to mitigate attenuation bias (such as using multi-year averages of father's income or IV), this will yield very nearly bias-free estimates. This approach is confirmed by Haider and Solon. Using data from the Health and Retirement Study the authors pursue a formal approach like that described above and estimate that $\alpha_S(t_S) \cong \alpha_F(t_F) \cong 1$ when fathers and sons are both observed around age 40.

If data for either generation is not available at midlife, a significantly less desirable rule of thumb is to choose data for each generation drawn from a similar point in the lifecycle. For instance, often data for the younger generation is not available much beyond age 30. In such cases, observing father's earnings early in the lifecycle would reduce the lifecycle

[16] This is not to mention the worries of bias should the parameterization of the ageearnings profile be incorrect or if the data used to estimate the age-earnings profile is not drawn from the same population as that used in the primary analysis as would have to be the case in any study of Britain using NCDS data.

bias.[17] However, the bias will be eliminated if and only if $\alpha_S(t_S)=\alpha_F(t_F)$ for the chosen age. Because α_I increases over the lifecycle and the average α_i is unity, we know $\alpha_S(t_S)$ must be relatively close to $\alpha F(t_F)$ sometime around midlife, but this is not to say that $\alpha_S(t_S) \cong \alpha_F(t_F)$ for all $t_S=t_F$. Indeed, the tremendous change in educational attainment experienced over the last 40 years nearly guarantees that age earnings profiles have changed between generations. Reference to Fig. 1 confirms that as we move away from midlife, the 'common age' rule will be less and less effective. One might suggest observing fathers at a slightly younger (older) age than sons when data are drawn before (after) midlife. Even though this second-best rule of thumb has obvious limitations, it would almost certainly be preferable to observe both fathers and sons at age 30 rather than sons at 30 and fathers at 55. The latter approach would more or less maximize the degree of lifecycle bias.

In conclusion, as Jenkins notes it is impossible to estimate β, the elasticity of total lifetime earnings, in the absence of a lifetime of data for both father and son. We can, however, address lifecycle bias when estimating the structural parameter g, the elasticity of lifetime permanent earnings. Formal correction of lifecycle bias requires extraordinary assumptions about data and are likely to produce large standard errors due to a multi-stage estimation process. As a result, informal methods may be as useful. Application of economic theory leads us to prefer estimates based on mid-life observation of both fathers and sons (when $\alpha_S(t_S)=\alpha_F(t_F)\cong 1$).

5. Conclusion

An examination of intergenerational earnings persistence estimates shows a strong, negative relationship between estimated persistence and the age at which father's earnings are observed. In total, 20 percent of the variance among studies employing similar estimation methodologies can be explained by differences in the age of fathers at observation. Whatever the cause, the strong dependence of persistence estimates on the father's age (or year of observation) alters our understanding of cross-country earnings persistence comparisons. For instance, estimates of persistence in Finland (Österbacka 2001) and Germany (Couch and Dunn 1997) which may otherwise seem extremely low appear typical once the relatively old age of fathers in these studies is considered.

Using the model of Jenkins (1987) two possible explanations for this regularity are explored: a) errors-in-variables attenuation bias may have increased over time producing yeareffects that appear like age-effects and b) lifecycle increases in the variance of the

[17] This is, of course, not the first work to recommend using intergenerational earnings data from similar points in the lifecycle. One of the earliest examples of this recommendation is found in Atkinson et al. (1983). However, none of these earlier works suggests the lifecycle bias explored in this paper as the reason for the suggestion. For example, Atkinson et al. cite changes in *mean* earnings over the lifecycle (p. 7) and attenuation bias (p.6) as reasons to prefer a common observation age. These concerns are addressed in the modern literature by including age controls and employing IV estimation respectively. This paper shows that while these strategies mitigate the important issues raised by Atkinson et al., they do not account for bias related to lifecycle patterns in earnings *variance*.

permanent component of earnings (predicted by human capital theory) produce a lifecycle bias. The data challenge the former explanation as researchers have not found time trends in intergenerational earnings persistence and transitory earnings variance has actually diminished in importance over time. Consistent with the latter hypothesis, within country study also finds that intergenerational earnings persistence estimates decrease with father's age and increase with son's age. This pattern is found in the US, Canada, and Germany even though the well-documented increase in transitory earnings variance has been experienced to very different degrees in the three countries.

The penultimate section of the paper discusses strategies for mitigating lifecycle bias when estimating the intergenerational persistence of lifetime earnings. To address the issue formally, measures of lifetime earnings are required. It is possible to estimate lifetime earnings with less than a full earnings history and then use these estimates to correct persistence estimates. However, this approach requires more data than is usually available and/or strong parametric assumptions. Furthermore, the multiple steps involved in the estimation process will likely create large standard errors in all but the largest samples. Given these limitations with the formal correction, simple rules of thumb may be just as useful. By observing both fathers and sons near midlife the bias is likely reduced.

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Appendix A

A.1. National longitudinal survey

Father's wage and salary income is recorded in 1966, 1967, 1969, and 1971 for the year prior to the survey. Fathers are restricted to be no older than 55 in 1966 to ensure that a selection bias is not introduced as older fathers retire in later periods. Positive earnings must be reported to be included in the sample.

The sons drawn from the Young Men Cohort are restricted to be no older than 18 in 1966 to avoid oversampling of sons who live at home after high school. Son's wage and salary income from the previous year is reported in 1971, 1973, 1975, 1976, 1978, 1980, and 1981. Given the young age of the respondents, the 1971 and 1973 data are not used.

To be included in the sample, the son must report positive earnings. In cases in which more than one son is available from a given household, only the oldest son in the sample is used. Note that this may not be the oldest son in the family since an older son may not have been included in the survey or the sample. The sample sizes range from 270 to 367 depending on the observation years of fathers and sons.

A.2. Intergenerational income data

The construction of the IID from Canadian tax files is described in detail in Corak and Heisz (1999). The sample studies families with children ages 16–19 in 1982. A one–in–ten sample was taken from the full data set and then, from this sample, the oldest available son for each family was selected. (Note, the oldest available son may or may not be the oldest son in the family.) This resulted in 56,141 father-son pairs. The data was then limited to those fathers born between 1932 and 1942 (inclusive) in order to avoid attrition bias since father's labor income is recorded from 1978 to 1992. Son's labor income is recorded from 1991 to 1998. The sample includes only observations with positive earnings reports.

Through an examination of the mean and variance of reported incomes, several coding irregularities were found. It appears that a significant number of observations in 1978–1982 were assigned a value of $1 when, in other years, they would have been reported as $0. Similarly, in 1996, a significant number of observations were assigned earnings of $2. It was not possible to determine why the data included these anomalies. 'Positive earnings reports' refer to incomes greater than $1 in 1978–1982 and greater than $2 in 1996.

A.3. Panel study of income dynamics

Sons, 9 to 17 years old at the time of the initial 1968 PSID survey, are observed from 1983 to 1992. The exclusion of younger sons ensures that the observations of son's income is not overly affected by non-representative observations at the beginning of the career. Exclusion of older sons avoids over-representation of sons who live with their parents beyond high school. Since head labor income is used to measure earnings, the son must be the head of household in the observation period in question to be included in the sample. Non-positive earnings reports are excluded. In families in which there is more than one son which fits these restrictions, the sample includes only the oldest available son.[18]

'Fathers' in the sample are the male heads of the households in which the sons lived in 1968. They are observed in the years 1967 to 1981. Fathers are eliminated from the sample if their age does not fall between 30 and 46 (inclusive) in 1967. Inclusion of older fathers who will likely retire during the observation period would introduce a sampling bias. Again, fathers must be heads of household in the observation period in question and report

[18] The study was replicated using the sample of all sons. The results do not change substantially with this alternative sample definition.

positive earnings. The resulting sample sizes range from 199 to 260 depending on the observation years of fathers and sons.

A.4. German socioeconomic panel

The GSOEP is a household panel survey with design and topic coverage similar to those of the PSID. The data for this study are drawn from the West German sample which was collected from 1984 through 2001. Sons, 13 to 17 years old in 1984 are only 30 to 34 in 2001, the final year for which data is available. And so only this one year of earnings data is collected for sons. Exclusion of older sons avoids over-representation of sons who live with their parents beyond high school. Non-positive earnings reports are excluded. Given the very small sample available, the sample was not restricted to only one son per family.

Male heads of household in 1984 form the population of 'fathers' in the sample. They are observed from 1984 to 1995. Fathers are eliminated from the sample if their age does not fall between 30 and 46 (inclusive) in 1984. Inclusion of older fathers who will likely retire during the observation period would introduce a sampling bias. Positive earnings in a given year are required to be included in the sample. The sample sizes range from 97 to 127 depending on the observation years of fathers and sons.

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