INTERGENERATIONAL LABOR SUPPLY

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This paper investigates whether the work behavior of children is strongly related to that of their parents. The motivation for studying this question is as an attempt to explain the wide variation of the annual hours of work of individuals often observed by researchers. Since much of this variation is attributed to individual tastes for work, much of which are thought to be attributable to family socialization, analysis of intergenerational correlations of labor supply behavior will provide insights into the source of individual tastes for work. This research uses intergenerational data from the Panel Study of Income Dynamics to examine intergenerational associations in work behavior. Overall, positive associations were found in the work behavior of fathers and sons. The intergenerational correlation in annual hours of work was found to be .09, almost all attributable to correlations in average weekly hours. In the analysis of mother-daughter pairs, few positive correlations in work behavior were found, although this result is most likely caused by an overrepresentation of working daughters in this sample. Overall, the intergenerational associations in labor supply behavior found in this paper were not as large as might have been supposed if the majority of one’s tastes for work were derived from one’s socialization within the family unit.

I. INTRODUCTION

* I would like to thank those whose assistance has been especially helpful to me in my work on this paper. I am greatly indebted to John Pencavel for his invaluable support, advice and guidance through the entire process of writing this paper. Ninny Khor was also a great support in her assistance with statistical programming.
When the average individual thinks of the typical work year, most think of the "standard" 5 day, 40 hour week, for 50 weeks a year. The general conception, then, is that the majority of individuals work 2000 hours each year, with little deviation. Yet, in reality, this "standard" work year is far from the norm. Researchers consistently have found a wide variation in hours worked across individuals in a given year as well as for some individual across years. Even when controlling for hourly wages and non-wage income, a large variation in hours of work persists. Researchers have tried to explain the remaining variation through observable differences across individuals that might affect one’s work opportunities, including non-wage income, educational attainment, age, marital status, race, region, conditions in the local labor market, union status, disability, etc. All of these variables have been found to have important effects on the labor supply of individuals. Yet even with these controls, only between 9 and 13% of the variation across individuals is explained (Pencavel 86)

If all these observable characteristics account for such a strikingly small percentage of the variation in work hours across individuals, what is driving the observed differences in labor supply? The main conjecture has been that preferences for work, unobservable to the researcher and not highly correlated with observable characteristics, are the primary explanation. This hypothesis is intuitively satisfying, as everyone knows individuals who simply enjoy working more than others. The question, then, is how these preferences for work are formed, and if measurable, how much of the variation in hours of work do they explain?

One important hypothesis of the formation of labor supply preferences is that socialization within the family from an early age is an important determinant of one’s own tastes for work. Anecdotally, sons of “hard-working” fathers tend to work longer hours themselves; daughters of mothers who chose to work are more likely to work as well. Children are likely to model the work behavior of their parent of the same gender. This means sons are likely to mimic the behavior of their fathers, daughters that of their mothers. This idea is similar to that of children choosing to pursue occupations similar to those of their parents. Therefore, the notion is that some unobservable taste for work is passed on from parents to children through family socialization and modeling of behavior.
How does this theory of family socialization of tastes for work relate to the problem of trying to explain variation in annual hours of work? The thought is, that, if the most significant source of preferences for work is cultivated within the family, these tastes could explain a significant portion of the remaining variation in hours worked. If one were to look at permanent hours of work, removing transitory shocks, this theory would imply that, for a large sample of individuals, the annual hours of work of sons would be highly correlated with the annual hours of work of their fathers, and annual hours of work of daughters with that of their mother. It is important to note that the labor supply decisions of one’s parents are unlikely to explain all of the remaining variation across individuals, especially as many have found a large amount of variation in labor supply amongst siblings (Solon 1991), yet a quantification of the effect of parent work behavior on that of their children would be useful in further examination of labor supply.

For a theory that is so widely assumed to be true, little empirical work has been done to test its validity. Economists have done little research to test how much parental preferences for work influence their children’s labor supply decisions. Many researchers have examined other dimensions of the work experience across generations, studying everything from earnings correlations [Solon (1992), Altonji (1991), Atkinson (1983)], to union status (Blanden & Machin 2003), to occupational choice, yet few have looked at intergenerational correlations in labor supply variables.

An important exception is the work by Altonji and Dunn (1991) using National Longitudinal Survey panel data to examine family links in labor supply. They examined the relationships between siblings, individuals related by marriage, but their most relevant work is that looking at intergenerational correlations between parents and children, with sample sizes ranging from 618 to 1118. In their analysis of fathers and sons, they found correlations of log annual hours to range between .058 and .190, and for mothers and daughters to be between .063 and .128. They also examined various dimensions of the labor supply experience, including weeks worked, average weekly hours, and weeks of unemployment. The strongest relationship found was in log weeks worked between father-son pairs (.081-.326), and mother-daughter pairs (.108-.522). These correlations indicate a fair degree of intergenerational correlation in labor supply, and as such, merit further study. Additionally, Altonji and Dunn did not look at the labor
force participation decision, an especially important source of variation in hours worked for females (Duncan 1977). As such, some important dimensions of intergenerational relationships in work behavior have yet to be researched at all.

This paper will expand upon the work of Altonji and Dunn (1991) in examining intergenerational correlations of labor supply behavior. The analysis in this paper will focus on father-son and mother-daughter pairs, utilizing the theory of gender role modeling of work preferences. The analysis will examine decisions to work for pay across generations, as well as intergenerational correlations in annual hours, weeks worked, average weekly hours and unemployment amongst those choosing to be in the labor force. Sections II, III and IV describe this new analysis of intergenerational correlations of work behavior, using panel data from the University of Michigan Panel Study of Income Dynamics (PSID), a dataset particularly well suited to this sort of analysis. Overall, positive associations in the work behavior of fathers and sons were found, with the intergenerational correlation in the annual hours of fathers and sons equal to .09. For mothers and daughters, no positive associations in work behavior were found.

While the primary motivation for studying intergenerational labor supply correlations is to explore the question of whether family socialization is an important factor in forming individual work preferences, it is important to note a secondary motivation for exploring this relationship. In work by Solon (1992), intergenerational associations in long-run earnings were found to be 0.4, indicating a low degree of income mobility, an issue of great interest to public policy analysts and social scientists. Yet little work has been done to study the mechanisms by which earnings are transmitted. Labor earnings are the product of hourly wage times hours worked, and as such, correlations in hours worked could be an important explanation for the large intergenerational earnings correlations. A study by Duncan & Morgan (1977) using the PSID, decomposed the variance in logarithmic change in labor earnings of individuals into the variance in logarithmic change in work hours, hourly wages, and a covariance term. For males, while the wage rate variance dominated that of work hours, the change in hours term was surprisingly large, indicating that differences in hours worked play an important role in explaining differences in male labor earnings. For females, the variance in annual hours dominated that of wages in its contribution to the variance in labor income. In both cases,
the variance in hours was an important component of earnings variance, indicating that a study of the degree of intergenerational correlation in labor supply variables could shed light on mechanisms important to earnings mobility.

**II. DATA DESCRIPTION**

The analysis in this paper uses intergenerational data from the Panel Study of Income Dynamics (PSID), a nationally representative longitudinal survey carried out by the University of Michigan Survey Research Center annually since 1968. The survey tracks demographic and economic behavior not only of adults from the original sample, but also of children as they transition into adulthood and form their own families. As such, the survey provides individually reported data on labor supply for both parents and children over time, allowing the work behavior of children to be related to that of their parents. Low overall attrition rates in the survey combined with the wealth of information available makes these data especially appealing for intergenerational analysis.

This study focuses on father-son and mother-daughter associations in labor force participation, annual hours of work, weeks worked, and average weekly hours, in addition to unemployment correlations for fathers and sons. The data in this study come from five years of observations for both adults and children. Adults are observed from 1968-1972, children from 1989-1993. Each set of observations contains data both at the family and individual level. The sample contains 348 father-son and 378 mother-daughter pairs from the SRC cross-sectional sample of the PSID. This sample was an equal probability sample of households from the 48 contiguous states, designed to yield 3,000 family interviews. The other sample, from the Survey of Economic Opportunity (SEO), a specially designated nationally representative sample of 2,000 low-income households in 1968, has been excluded from the analysis of this paper. This sample over represents low-income households and, if low-income household’s tastes for work were systematically different from the rest of households, an overrepresentation of low-income households would introduce systematic biases of one kind or another.
Sons in this sample are children from the original 1968 PSID households whose fathers were head of household in each of the five years between 1968-1972. Only those sons responding to the survey in each of the five years of interest, 1989-1993, were included in the sample. As 99% of sons worked in these years, only sons working positive annual years in each year from 1989-1993 were included in the sample. Additionally, sons were restricted to be between the ages of 25 and 50 in 1989 to try to observe fathers and sons at the same age. Finally, only sons who had moved away from home and were head of a household in all five years from 1989-1993 were included. In the case that more than two sons in a family met the requisite criteria, only the oldest was retained in the sample in order to observe the sons most likely to be near his long-run labor supply levels.

Fathers in this sample were heads of household where sons resided during 1968. Fathers were not necessarily the natural fathers of their sons, as the focus of this study is about socialization within families, something that occurs regardless of the blood relation of a father to his son. Only fathers who responded to the survey in each year from 1968-1972 and were head of household in these years were retained. Finally, fathers were constrained to be between the ages of 25 and 50 in 1968, as stated before, in order to observe fathers and sons at similar ages.

Although sons were constrained to work positive annual hours in a given year, fathers were not. Therefore, fathers working no annual hours in a given year were retained in the sample. The reasoning is that sons are constrained in that at a minimum

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1 Any married male is classified as head of household whether he works or not, eliminating the possibility of bias for males who work versus those choosing to stay at home

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
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<tr>
<td>Son's Age in 1991</td>
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<td>47</td>
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<td>96</td>
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</table>
they can work zero hours, even if they would prefer to work negative annual hours. This censoring of the sons annual hours means that including sons working zero hours could represent sons wanting to work either zero or fewer hours. However, since these sons must work zero hours, inclusion of these sons can cause the association between annual hours of work of father and sons to be misestimated. This censoring problem was pointed out in Killingsworth (1983). Censoring the annual hours of work of the father, however, does not cause any bias in the estimation of the relationship between annual hours of fathers and sons, and as such, the inclusion of these fathers is acceptable.

Table 1 presents descriptive statistics for fathers and sons. Although the sample tried to select fathers and sons at close to the same age, the mean age of sons is 35, whereas that of father is 42. Table 1 also reports that the mean annual hours of work of sons is higher than that of fathers with smaller variance, but this result most certainly can be explained by the inclusion of fathers working zero annual hours in the sample.

Daughters and mothers for the main sample were selected in an almost identical manner as sons and fathers were. Both had to respond to the survey in the appropriate years, and be between the ages of 25 and 50 in the first year each was observed for this analysis. An additional criterion was that daughters must be head of a household or a wife in 1989, excluding those daughters still living at home or residing with a sibling in 1989. This restriction was imposed to avoid over representing daughters leaving the home at late ages. Mothers could either be a head of household or wife in a household where a daughter resided during 1968.

Table 2 presents descriptive statistics for important characteristics of mothers and daughters in the main sample. Despite attempts to observe mothers and daughters at the

<table>
<thead>
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<th>Table 2-Sample Characteristics of Mothers and Daughters</th>
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<tr>
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<tr>
<td>Daughter’s Age in 1991</td>
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<tr>
<td>Daughter’s Annual Weeks Worked in 1991</td>
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<tr>
<td>Daughter’s Average Hours per Week in 1991</td>
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<tr>
<td>Mother’s Annual Hours of Work in 1970</td>
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<tr>
<td>Mother’s Age in 1970</td>
</tr>
<tr>
<td>Mother’s Weeks Worked in 1970</td>
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<tr>
<td>Mother’s Average Hours per Week in 1970</td>
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</tbody>
</table>

N=378
same age, the mean age for mothers is 40 as opposed to 35 for daughters, a five-year difference. Daughters are more likely to work than their mothers, with 84% working in 1989 versus 47% of mothers in 1968. Working, in this paper is defined as working at least one hour per year. As an additional measure of the decision to work, 93% of daughters chose to work at any point during the five years they were observed, compared with 71% of mothers. Annual hours of work by daughters are more than twice that of the mothers and have a higher standard deviation. These findings are consistent with the increasing labor supply of women in more recent cohorts observed by many researchers (Killingsworth & Heckman 1986). While not reported in Table 2, it is important to note some descriptive statistics for observed characteristics of mothers and daughters, such as marital status, educational attainment, and the presence of young children. Overall, daughters in the sample are less likely to be married than their mothers, and have a higher level of educational attainment, all consistent with observed demographic trends.

The main sample mother-daughter pairs were used to examine the labor force participation probabilities. When examining correlations of annual hours of work, weeks worked, average weekly hours and unemployment, the sample was further restricted to those daughters who worked positive hours every year from 1989-1993. No additional restrictions were placed on the parents for this sample. This selection reduced the sample of mother-daughters to 274 pairs. By limiting the sample in this manner, when studying correlations in hours of work, one is not examining overall correlations in the population, but rather asking ‘Among working children, what is the link between the hours of work of one generation (parents) and the hours of work of the next generation (children)?’ Posing this question avoids the problem of censoring the left hand side variable at zero; an especially important issue for estimation of mother-daughter regressions, as in any given year approximately 15% of daughters did not work (Killingsworth 1983). Table 3

<table>
<thead>
<tr>
<th></th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
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<td>4140</td>
</tr>
<tr>
<td>Daughter’s Age in 1991</td>
<td>35.51</td>
<td>4.71</td>
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<td>46</td>
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<td>Daughter’s Annual Weeks Worked in 1991</td>
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<td>9.46</td>
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<td>Daughter’s Average Hours per Week in 1991</td>
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<td>98</td>
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<tr>
<td>Mother’s Annual Hours of Work in 1970</td>
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<td>881.92</td>
<td>0</td>
<td>3770</td>
</tr>
<tr>
<td>Mother’s Age in 1970</td>
<td>40.10</td>
<td>6.41</td>
<td>27</td>
<td>52</td>
</tr>
<tr>
<td>Mother’s Weeks Worked in 1970</td>
<td>20.98</td>
<td>21.98</td>
<td>0</td>
<td>52</td>
</tr>
<tr>
<td>Mother’s Average Hours per Week in 1970</td>
<td>19.01</td>
<td>19.14</td>
<td>0</td>
<td>73</td>
</tr>
</tbody>
</table>

N=274
presents modified summary statistics for the reduced samples. Besides hours of work for the children, the underlying population characteristics for the individuals in this sample are essentially identical to those in the original sample.

III. Econometric Models

This section presents the models used to relate the labor supply decisions of children to those of their parents, and is divided into three parts, each examining a different set of econometric models estimated in analysis of this paper. The first part presents models used in evaluating intergenerational correlations in annual hours of work, annual weeks worked, and average weekly hours. The second presents those used to predict the probability of labor force participation of daughters based on the participation of their mothers, and the third those used in the prediction of unemployment probabilities of sons conditional on the unemployment of their fathers.

A. Annual Hours of Work

The hypothesis of this paper is that tastes for work are transmitted between generations. To test this theory, one would conjecture that annual hours of work of a child and that of his or her parent would be strongly correlated. Let $h_{ic}$ represent the long-run component of annual hours (the “permanent” component of annual hours) worked by a child in family $i$, and $h_{ip}$ represent the same for the parent in family $i$. Assume that the variation of $h_{ic}$ and $h_{ip}$ are the same. Then, if $h_{ic}$ and $h_{ip}$ were observed from a random sample of families, the true population correlation $\bar{n}$ could be estimated by applying least squares to the regression equation

\[
(1) \quad h_{ic} = k + \bar{n}h_{ip} + e_i.
\]

where $k$ is the intercept, as sons are restricted to working positive annual hours. A major problem in using this equation to estimate the true population correlation between $h_{ic}$ and $h_{ip}$ is that the researcher does not observe the permanent hours worked by each individual. Instead, short-run measures of annual hours are used as proxies for the true long-run values. The relation of long and short run measures in period $t$ for a child in family $i$ can be expressed as

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2 Annual weeks worked and average weekly hours models were estimated similarly to those used for annual hours.
\( h_{ict} = h_{ic} + v_{ict} \)

where \( v_{ict} \) is the transitory variation of \( h_{ict} \) from \( h_{ic} \). The fluctuation of \( h_{ict} \) could be due to actual variation in a given year from the permanent annual hours worked by an individual, or could represent measurement error. A similar equation estimated for a parent in family \( i \) at time \( s \) is

\( h_{ips} = h_{ip} + v_{ips} \)

Assuming that the transitory components are uncorrelated with each other and the long run variables around which they fluctuate, using \( h_{ict} \) and \( h_{ips} \) in place of \( h_{ic} \) and \( h_{ip} \) in equation (1) will tend to underestimate \( \bar{n} \) in relation to its true population value. Solon (1992) made this point in his work on intergenerational income correlations. The downward bias of \( \bar{n} \) in the study of intergenerational annual hours correlations is even more of a concern than in the study of income correlations, as studies of the accuracy of PSID data found that annual hours of work data tend to have more measurement error than that of annual earnings (Bound et al., 1989). Any downward bias found in Solon’s work would therefore be more severe in the analysis of this paper. As such, the downward bias in annual hours correlations could be quite significant.

One source of fluctuation in annual hours of work is age, which is known to have a distinct relation with annual hours of work. As such, because the analysis in this paper uses short-run data, equations (2) and (3) are expanded to include the age of both children and parents. Additionally, in some estimations, the child’s annual hours specification will include a set of controls for observable characteristics shown to have important effects on annual hours of work. Therefore, for children, equation (2) is extended to

\( h_{ict} = h_{ic} + \hat{a}_c + \hat{a}_{ict} + \hat{a}A^2_{ict} + v_{ict} \)

where \( A_{ict} \) is the age of a child in family \( i \) at time \( t \). The model for parent’s annual hours is extended to

\( h_{ips} = h_{ip} + \hat{a}_p + \hat{a}A_{ips} + \hat{a}A^2_{ips} + v_{ips} \)

where \( A_{ips} \) is the parent’s age in year \( s \). Solving equations (4) and (5) for \( h_{ic} \) and \( h_{ip} \) and substituting the resulting expressions into equation (1) gives

\( h_{ict} = k + (\hat{a}_c - \bar{n}\hat{a}_p) + \bar{n}h_{ips} + \hat{a}_{ict} + \bar{n}\hat{a}A_{ict} + \bar{n}\hat{a}A^2_{ict} - \bar{n}\hat{a}A_{ips} + \bar{n}\hat{a}A^2_{ips} + e_i + v_{ict} - \bar{v}_{ips} \)

While this model addresses the short-run nature of the data used in the analysis of this paper, its use alone to estimate intergenerational correlations in tastes for work would
be naïve, as labor supply decisions represent a combination of preferences and constraints. Individuals make decisions about much to work in a given year based not only on their tastes for work, but also on constraints they face (Killingsworth 1983). The most important of these is the budget constraint. One’s hourly wage as well as one’s level of non-wage income play an important role in determining the level of one’s labor supply. At a higher wage, one might choose to work more or less hours than they would at a lower wage. Thus, to analyze the effect of preferences, any model of labor supply should include these constraints. As such, equation (1) is more correctly represented by

\[ h_{ic} = \bar{h}_{ip} + \hat{a}w_{ic} + \hat{a}y_{ic} + e_i, \]

where \( w_{ic} \) represents the permanent average hourly wage of a child in family \( i \), and \( y_{ic} \) represents before-tax non-wage income. Use of this model more accurately assesses the connection of tastes for work between generations. It should be noted that, in many studies of labor supply, average hourly wage and non-wage income explain a relatively small fraction of the variance in labor supply across individuals, but their effects are significant, and as such, should be included in the specification of a model of labor supply.

Some might argue that individuals face other constraints in their labor supply decision besides budgetary ones. For example, at a constant wage rate, an individual might prefer to work 41 hours a week but the job that best maximizes their utility might only allow them to work 40. If present, these sorts of constraints could affect the results estimated in equation (7), though it is difficult to predict what effect these constraints have. While it is true that such constraints could be important, they are unobservable to a researcher, and as such, cannot be included in equation (7). However, intuition implies that tastes for work would explain more of the variation in annual hours of work than labor supply constraints above and beyond those of budget constraint. Therefore it seems appropriate to say that a significant portion of any correlation found between the annual hours of parents and their children derive from preferences.

To incorporate the budgetary constraints of children into the model used in this analysis, equation (6) is refined, substituting the expressions for \( h_{ic} \) and \( h_{ip} \) derived from equations (4) and (5) into equation (7), resulting in
\begin{equation}
(8) \quad h_{ict} = (\hat{a}_c - \hat{\bar{a}} d_p) + \hat{\bar{n}} h_{ips} + \hat{\bar{a}} w_{ict} + \hat{\bar{a}} y_{ict} + \hat{\bar{a}} A_{ict} + \hat{\bar{a}} A_{ict}^2 - \hat{\bar{n}} \hat{\bar{d}} A_{ips} - \hat{\bar{n}} \hat{\bar{d}} A_{ips}^2 + e_i + v_{ict} + \hat{\bar{n}} v_{ips}.
\end{equation}

where \( w_{ict} \) and \( y_{ict} \) are observed for the child at time \( t \) and the rest of the parameters are the same as in equation (6).

Equation (8) expresses a child’s observed annual hours in year \( t \) as a regression function of parent’s observed annual hours in year \( s \), budget constraint controls for the child in year \( t \), age controls for both the parent and children, and controls for observed characteristics of the children. As stated, the estimation of \( \bar{n} \) is subject to a downward bias due to the use of short-run proxies for long-run annual hours of work.

This paper employs two main methods to address the likely downward bias of \( \bar{n} \). The first is to use more long-term measures of annual hours. For these equations, the status of the parents and children are averaged over \( T \) years, modifying equation (7) to

\begin{equation}
(9) \quad \hat{\bar{a}}_{ict} = (\hat{a}_c - \hat{\bar{a}} d_p) + \hat{\bar{n}} h_{ips} + \hat{\bar{a}} w_{ict} + \hat{\bar{a}} y_{ict} + \hat{\bar{a}} A_{ict} + \hat{\bar{a}} A_{ict}^2 - \hat{\bar{n}} \hat{\bar{d}} A_{ips} - \hat{\bar{n}} \hat{\bar{d}} A_{ips}^2 + e_i + v_{ict} - \hat{\bar{n}} v_{ips}.
\end{equation}

where any variable \( \hat{\bar{a}}_{ips} = 1/T \ast \hat{\bar{O}} u_{ips}^3 \). While this estimation does not completely remove the downward bias of \( \bar{n} \), it does decrease its magnitude. By using multi-year averages of annual hours, one is reducing the likelihood of observing an individual during an uncharacteristic year of their work behavior and therefore can better estimate the true correlation of tastes for work between generations.

The second method applied to address the error-in-variables problem present in equation (8) is that of instrumental variables. For this paper, a set of several instruments for a parent’s annual hours of work was used. The instruments chosen vary across gender, as different variables are more important in determining annual hours of work for males and females.

For males, the set of instrument used for fathers include: average hourly wage, non-wage income, whether the father was disabled, a father’s union status, whether the father was unemployed during the previous year, whether the father worked a second job, and a father’s education. While it is not clear that the father’s average hourly wage, non-wage income and education do not belong in the original specification of equation (8), because this paper aims to construct a model of the intergenerational correlation in annual
hours worked, these were not included in the original model as to capture the full effect of annual hours of work of the father on the son. The use of average hourly wage, non-wage income and education as instruments is a natural selection if not included in the original model for annual hours of work of the son, as these three variables have been shown to have important effects in the determination of annual hours of work of an individual (Pencavel 1986). The selection of the other four variables as instruments was based on the work of Duncan and Morgan (1977), who found these variables have important effects in determining the change in an individual’s annual hours of work across years. For mothers, average hourly wage, non-wage income, education, and the presence of children either between the ages of zero and six or six and eighteen were the instruments selected. The reasons for choosing these variables as instruments for annual hours of work of mothers were similar to the reasons for selection of instruments for the fathers.

These instruments only will consistently estimate $\hat{n}$ if they do not belong in equation (8) and they are perfectly correlated with the annual hours of work of the parent. Considering that regressions using these variables as explanatory variables for annual hours of work only typically explain up to twenty percent of variation in annual hours of work, the use of instrumental variables is unlikely to consistently estimate $\hat{n}$. The direction of the inconsistency, however, is unclear. As opposed to the work by Solon (1992) on intergenerational income correlations, in which the inconsistency was most likely upward, the relation of the set of instruments used in this analysis to annual hours of work of sons is unclear. This means that the inconsistency of estimation in instrumental variables could be upward, downward, or even consistent.

**B. Labor Force Participation**

For mothers and daughters, an important dimension of labor supply behavior is the decision of whether to work for pay\(^4\). Therefore, any research of intergenerational work behavior of females without analysis of the probability of labor force participation of daughters would be incomplete.

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\(^3\) Qualitative characteristics were formed in the same way as quantitative variables in multi-year analysis.

\(^4\) For fathers and sons, 98-99% of each generation worked, leaving almost no variation with which to study labor force participation, and as such, this analysis is not included in this paper.
To analyze probability of a whether a daughter will participate in the labor force, one would like to estimate the linear probability model

\[
(10) \quad \text{prob}(I(D)=1|I(M)) = a + b[\text{prob}(I(M)=1)] + e
\]

where \(\text{prob}(I(D)=1)\) is the probability of a daughter working for pay, and \(\text{prob}(I(M)=1)\) is the probability for mothers. In this equation, \(b\) measures the conditional probability of a daughter working, in which \(\text{prob}(I(D)=1|I(M))\) is the probability of a daughter working, conditional on a participation of her mother in the labor force. Participation in the labor force is defined as working positive hours in a given year. Of course, in any given year an individual either works or does not, but the interpretation of \(b\) is still the same.

One problem with estimation of equation (10) is that it does not fully account for factors important in the decision of daughters to work or not. As discussed by Killingsworth and Heckman (1986), factors such as non-wage income, marital status, age, education, the presence of young children, and other factors have important effects on the decision of females to work. While one could argue that the decision of a daughter to get married or have children is strongly influenced by the decisions modeled by their mother, this study aims to look purely at the intergenerational associations in tastes for work. As such, equation (10) is extended to

\[
(11) \quad \text{prob}(I(D)=1|I(M), X) = a + b[\text{prob}(I(M)=1)] + c.X + e
\]

where \(X\) is a vector of right hand side variables including age variables for mothers and daughters as well as marital status, education, and presence of young children for daughters. While equation (11) does not control for differences in the opportunities presented to daughters choosing to work or not, it does a better job estimating the true link of preferences between mothers and daughters than a model not taking these factors into account.

While equation (11) is a better specification of the probability of labor force participation of daughters, the underlying assumption is that the long-run labor force status of mothers and daughters is the variable used in the estimation of equation (11). Ideally, a researcher would have access to the entire work history for mothers and daughters, and from such information could ascertain whether on average each individual participated in the labor force. Of course, in practice, one does not have this sort of
information. Instead, one must use the observed labor force status of an individual in a given year as a short run proxy for long run behavior when estimating equation (11). The result is a likely bias in the estimation of \( b \). This errors-in-variables problem is slightly different than the one explored with annual hours of work, however. In this case, measurement error in a given year means that, for example, the researcher enters a zero for the labor force participation of an individual (meaning they did not work for pay) when in reality, in the long run, this individual did work for pay. The direction of the bias in the estimation of \( b \) is actually unclear in this case, although has been shown to most likely be downward. As such, through the rest of this paper, the estimation of \( b \) is assumed to be downwardly biased, though this assumption needs further analysis.

This paper uses a method similar to that presented in section IIIA to address the issue of bias in equation (11), using more long-run measures of labor force participation. Two types of long-run labor force status variables were created in this paper. For the first, the labor supply status for mothers and daughters was averaged over the five years each individual was observed. This variable can take on any multiple of one-fifth between zero and one, with participation in none of the years being equal to zero, 3 years to 0.6, and all years to 1.

The second long run labor force participation variable was constructed slightly differently. It takes on a value of one if an individual worked in any of the five years observed. This variable takes on a value of zero for individuals who never worked positive hours and a value of one if an individual worked positive hours at any time during the five years. The idea behind this conceptualization of labor force participation is that a woman not participating in the labor market in any of the five years observed has a strong preference towards not working. Her long-run labor supply status is most likely that of not participating. Conversely, a woman choosing to work in any of the five years has a stronger taste for work, and her absence from the labor market in a given year could likely be due to other constraints she faces, such as the presence of young children. Most of the results presented in the analysis of labor force participation links use this measure of labor force status. Overall, the use of either of these more long-run estimations of the labor market status of an individual reduces the chance of observing that individual
during a year not representative of their true participation in the labor market, reducing the likely bias in estimations of $b$.

As a final comment about the models used in the analysis of the probability of a daughter working for pay, equation (11) was specified by both a linear model and nonlinear probit functional form. The probability of participation in the labor force for daughters is restricted to be between zero and one, but in a linear estimation of equation (11), fitted values of this equation could actually be less than zero or greater than one. To force the fitted values of the dependent variable to be between zero and one, a probit functional form is used. When results for the probit estimations are reported in this paper, the derivatives of the probability of the labor force participation of a mother being equal to one is typically the quantity that will be reported.

C. Unemployment

As discussed in part A of this section, individuals often face constraints in the labor market beyond budgetary ones. One of the most obvious constraints is unemployment, which represents an individual in the labor force wanting to work more hours but unable to find an opportunity that allows him or her to work the hours they would like at an acceptable hourly wage rate. To better differentiate the contributions of constraints and preferences to intergenerational links in annual hours of work, this paper analyzes conditional probabilities of a son being unemployed$^5$.

Before discussing the models used in the analysis of unemployment links between generations, it is important to note that, while unemployment can be regarded as a constraint, to some degree, it also reflects a preference. When analyzing unemployment, the researcher does not observe the wages an individual is offered but does not take. A person who is unemployed may be offered jobs, just not at a wage rate acceptable to him or her. In that sense, unemployment is a preference as well as a constraint. This distinction is important, as one’s willingness to be unemployed might well be attributable to family socialization. A son whose father was unemployed might be more willing to be unemployed himself, whereas an individual whose family attached a negative stigma to unemployment might be willing to take lower wage jobs to avoid unemployment.

$^5$ Appropriate unemployment data for mothers was not available, and as such, analysis of intergenerational relationships in unemployment for females is excluded from this paper.
Therefore, a high probability of unemployment for a son conditional on the unemployment of his father can represent family socialization, just in a different way than in intergenerational correlations of annual hours of work.

The equations used to examine the probability of a son being unemployed are similar to those used in the analysis of labor force participation. However, it is difficult to imagine the permanent labor force status of an individual as being unemployed. Unemployment is, for the majority of individuals, a transitory state. Most spells of unemployment are short in duration. As an illustration, in 2003, 31% of unemployment spells were under five weeks, 31.7% were between five and fourteen weeks, and only 20.5% were over twenty seven weeks, the number typically used when defining an individual’s labor force status in a year as ‘unemployed’ (Bureau of Labor Statistics 2003). Additionally, very few individuals are unemployed over their entire working lifetime. Indeed, after several years of unemployment, one can imagine that most individuals would choose to exit the labor force.

Because being qualified as permanently unemployed does not make sense for the majority of working individuals, alternative measures of the state of unemployment for fathers and sons need to be constructed. Because at most 8% of observations for each generation in a given year were unemployed, analysis in this paper focuses on more long-term measures of unemployment as opposed to unemployment in a given year. The key variable used is a qualitative variable that takes on a value of 1 if an individual was ever unemployed during the five-year period they were observed. In this paper, unemployment is defined as having reported being unemployed for positive hours in a given year. The most general linear probability equation estimated was

\[
\text{prob}(\text{U(S)}=1|\text{U(F)}) = a + b[\text{prob}(\text{U(F)}=1)] + e
\]

where \text{prob}(\text{U(S)}=1) is the probability of a son being unemployed in a given year, and \text{prob}(\text{U(F)}=1) is the probability for the father. In this case, \(b\) measures the probability that a son was ever unemployed conditional on his father ever being unemployed. To control for observed factors in short-run measures of long-run variables, as with labor force participation, equation (12) is extended to include age controls for both fathers and sons, as well as control for non-wage income and other observed characteristics of sons.

\footnote{These values of the spells of unemployment were not seasonally adjusted.}
These are represented as a vector of variables ‘X,’ similar to equation (11). Finally, as with labor force participation, both linear and probit functional forms of equation (12) are estimated. The variable reported when the probit functional form was used is the derivative of the probability that a father was unemployed.

IV. EMPIRICAL RESULTS

This section presents the empirical results of the models outlined in Section III. The first two parts of this section present analysis of intergenerational correlations of males and females respectively. The first part is broken into two additional sections: annual hours analysis and unemployment for fathers and sons, while the second part presents labor force participation analysis followed by results of annual hours of work for mothers and daughters. The final part of this section discusses the implications of the results reported in the first two parts.

A. Fathers and Sons

1. ANNUAL HOURS OF WORK

Table 4 displays estimates of $\bar{n}$ from OLS estimation of equations (8) and (9). All average hourly earnings and non-wage income control variables are expressed in 1993 dollars as measured by the consumer price index. The first column of Table 4 shows the one-year estimates of $\bar{n}$ in each year $s = 1968, 1969…1972$, for $t=1991$. These results come from OLS estimation of equation (8), namely regressing a single year of annual hours of work of the son on a single year of the father. The $\bar{n}$s for all 25 one-year combinations of years of annual hours of work of fathers and sons (for example, 1968 for fathers with 1989 for sons, etc.) are presented in Appendix Table 1. For 1991, these estimates ranged from .047, when the year of the annual hours of fathers was 1971, to .099, using annual hours of the father in 1968. Two of these five coefficients were estimated to be statistically different from zero at the 10% level. For all 25 regressions, the one-year estimates ranged from .017 ($t=1989$, $s=1972$) to .122 ($t=1990$, $s=1970$), with 13 of the 25 statistically different from zero at the 10% level, and 4 at the 5% level. No single fact seems to explain this wide range of estimates, although those estimates using

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7 Any variable referred to as measured in 1968 refers to the interview year. Actual data are from the previous year, namely 1967, 1968…1971.
data where \( t=1989 \) for sons were smaller than most of the others, possibly due to a recession in the United States at that time.\(^8\) About half of the coefficients were between .07 and .09, and several of the coefficients falling in this range were statistically different from zero. All of these approximations of \( \hat{\beta} \) are expected to suffer from substantial downward bias.

To address problem of errors-in-variables, OLS is applied to equation (9), using multi-year averages for annual hours of work of both fathers and sons. These results are reported in the second and third columns of Table 4. The base year for annual hours of work of the son are reported in parentheses under the column headings, and the results for fathers were placed in the row of the median year in which they were observed (for example, results for father’s status from 1968-1970 are in the 1969 row). The range of these estimates is much smaller, ranging from .085-.092. All are statistically different from zero at the 5% level. Interestingly, while these multi-year estimates of the correlation in annual hours of work are larger than the majority of single year estimates as predicted, three of the one-year coefficients are actually larger than the multi-year estimates. The explanations for this phenomenon are unclear, but most likely represent

\(^8\) The mean and variance for 1989 annual hours of work were not the lowest of all the in which years sons were observed.
To further explore the robustness of the results presented in Table 4, other specifications of the correlation between father and son annual earnings were estimated by OLS. The first model was that of equation (6), dropping the budget controls and solely examining the intergenerational correlation of annual hours with age controls for the father and son. The second method added controls for the son’s education and marital status to equations (8) and (9). These results are presented in Table 5. Note that the second row of the table presents the same results as Table 4. The use of different controls had little effect on either the magnitude or the standard error of the coefficients estimated. For one-year measures, the addition of observable controls changes the estimate range from .047-.099 to .044 - .098. Results for the five-year averages of annual hours are very similar across specifications.

Overall, it seems convincing that the correlation between father’s annual hours of work and their son’s is approximately .09. This estimation is consistent with the work of Altonji and Dunn (1991), who report estimates of the correlation between annual hours of work of fathers and sons to be between .058 and .190. As a note, Altonji and Dunn used the log of annual hours of work in their research, and as such, removed all fathers working zero hours in a given year. These fathers are included in the analysis of this paper and this sampling difference could explain much of the deviation of the results.

Table 5

<table>
<thead>
<tr>
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<tbody>
<tr>
<td>Age Controls</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>0.105</td>
<td>0.068</td>
</tr>
<tr>
<td></td>
<td>0.045</td>
<td>0.047</td>
</tr>
<tr>
<td>Age and Budget Controls</td>
<td>0.099</td>
<td>0.063</td>
</tr>
<tr>
<td></td>
<td>0.045</td>
<td>0.047</td>
</tr>
<tr>
<td>Age, Budget and Observable Controls</td>
<td>0.098</td>
<td>0.063</td>
</tr>
<tr>
<td></td>
<td>0.045</td>
<td>0.047</td>
</tr>
</tbody>
</table>

N=348

Note (1): Standard error estimates are in parentheses.
Note (2): Observable controls include marital status and education of the son over the given time period.
To better understand the derivation of the correlation between annual hours of work of fathers and sons, this paper presents estimations of correlations for annual weeks worked and average weekly hours. The motivation for this analysis is that annual hours of work are calculated as the product of annual weeks worked and average weekly hours, and as such, both could be important in explaining the intergenerational links of annual hours of work. The methodology applied in studying these correlations is the same as that used for annual hours of work. Table 6 presents comparisons of the $\hat{n}$ coefficient as estimated by equations (8) and (9) for each labor supply measure. These equations are estimated with both one and five years of data. All of the equations had controls for age, budget, and observed personal characteristics (Note that the first row of Table 6 is equivalent to the third row of Table 5). Sample sizes for each set of regressions are in the far right column of each row.

One result that is immediately surprising is the negative correlation between annual weeks worked by the fathers and sons. The estimations of the correlation of annual weeks worked ranged from -.037 to .027. In analysis of fifteen regression estimations of $\hat{n}$ using different combinations of one-year data for fathers and sons

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9 For annual weeks worked and average weekly hours, data for fathers in 1968 and 1969 was reported in brackets, giving only a range of the weeks worked or average weekly hours for each individual. For these years, the average value for the category for fathers in the other three years was substituted for use in analysis. The unavailability of the true values of the variables in these years causes substantial measurement error, and accordingly, extreme downward bias in estimates of $\hat{n}$.
(excluding 1968 and 1969), eight were found to be negative, while none of the fifteen were statistically different from zero (See Appendix Table 2). One possible explanation is that sons who observe their fathers working a large number of weeks during a year might decide that their father’s quality of life was lower than it could have been from working fewer weeks, and decide to work fewer weeks themselves. This explanation does not seem very plausible, however, as the correlation between annual hours is statistically different from zero and positive. Also, for this explanation to hold, sons deciding to work fewer weeks in a year would have to work significantly fewer weeks each year, a result that does not hold up against the data presented. Finally, the work of Altonji and Dunn reported positive correlations in log weeks worked, contradicting the findings here. It is important to note that all negative estimations of \( \bar{n} \) in this paper were not statistically different from zero, meaning the distribution of the data made it difficult to accurately assess the correlation between the father’s and son’s weeks worked.

The high standard errors of the correlation coefficients for annual weeks worked indicate that the problem might result from the data. Examining a plot of weeks worked by sons in 1991 against those worked by fathers in 1970 shows strikingly little variation in the weeks worked by both fathers and sons. Splitting this same plot into four quadrants is even more revealing. The quadrants were assigned as follows: observations in the first quadrant were those for which both the father and son worked fewer than 30 weeks a year; the second and third quadrants contained observations for which at least one generation worked more than thirty weeks while the other generation worked fewer than thirty; and for observations in the fourth quadrant, both the father and son worked more than thirty weeks a year. Not surprisingly, the majority of the observations were bundled in the fourth quadrant. The first quadrant had no observations while the second and third quadrants had a few observations each. The significance of this result is that the observations in the second and third quadrants might have an important effect on the reported relationship between the annual weeks worked by fathers and sons.

To examine this hypothesis, father’s and son’s weeks worked were both restricted to be above 30 weeks in 1970 and 1991 respectively. This constraint dropped only 11 observations from the sample, leaving 339 father-son pairs. The regression estimation of equation (8) for weeks worked was then re-estimated, with \( t=1991 \) and \( s=1970 \). The
estimate of the correlation in weeks worked for this sample was -.032 with a SE of .062, compared to the -.037 with SE of .040 as estimated before. Clearly, restricting the observations of weeks worked to be greater than thirty weeks in both generations had little effect on the magnitude of \( \hat{n} \) estimated, but did increase the noise. It seems that such a large majority of fathers and sons worked 50 or more weeks a year, and that any attempts to estimate a relationship between the two generations weeks of work in this sample are extremely difficult.

In contrast to the results for weeks worked, the correlations between average weekly hours of fathers and sons are quite strong. In Table 6, estimates of \( \hat{n} \) for fathers and sons average hours of work in a single year ranged from .026 to .051. The estimate using five-year averages was .085, and statistically different from zero at the 10% level. These estimates are similar to those of Altonji and Dunn (1991), who estimated a correlation between .032 and .095. Since the first two years of data for fathers were bracketed, and estimates using these years of observations were likely to suffer from especially strong error-in variables biases, a new estimate of equation (8) was calculated, using a three-year average of the data, observing fathers from 1970-1972, and sons from 1991-1993. The result was an estimate of \( \hat{n} \) equal to .098 with a SE of .050, a result statistically different from zero at the 5% level. The magnitude of coefficient is similar to that found for annual hours of work of fathers and sons, suggesting that the main mechanism for transmitting annual hours could be through strong correlations in average weekly hours. Namely, tastes for annual hours of work are passed from fathers to sons in the form of the average hours one chooses to work in a given year as opposed to the number of weeks.

Finally, column seven of Table 6 presents results from instrumental variable estimations of equation (8) for annual hours of work, weeks worked and average weekly hours. For all estimates, \( s=1970 \) and \( t=1991 \), meaning fathers were observed in 1970 and sons in 1991. This estimation was done as an alternative method for addressing errors-in-variables problems. For annual hours of work, the instrumental variables estimation was .017 with a SE=.121, a result not statistically different from zero. In comparison with OLS results, IV estimation of \( \hat{n} \) greatly increased the “noise” with which the relationship was estimated, yet also weakened the “signal,” making it difficult to estimate a
relationship between annual hours of fathers and sons. In estimation of $\bar{n}$ for annual weeks worked and average weekly hours the results were similar, namely, that the standard error increased while the magnitude of the estimated coefficient decreased. Overall, as these results are on the lower end of estimates for each labor supply variable, it seems that this choice of instruments is downwardly inconsistent in its estimation of $\bar{n}$.

2. UNEMPLOYMENT

This part of section IV explores the probability of sons being unemployed given the unemployment experience of their fathers. Table 7 presents several estimates of $b$ from regression of equation (12) using different sets of controls. The variable used was whether the son or father was ever unemployed during the five years observed, having a value of 1 if the individual was unemployed, zero if not. Both OLS and probit results are reported, as when one has a discrete dependent variable, use of nonlinear analysis can be more appropriate. For the probit results, the quantity reported is the derivative of the probability of a father being unemployed. P-values for both OLS and probit analysis are reported in brackets to make the results comparable. The sample size is larger than that used to analyze annual hours, as sons could work zero annual hours in this sample.

Both the OLS and probit results display similar trends in their response to different sets of controls. As more controls are added, the estimated magnitude of $b$ for each type of model decreases and the P-values increases. Also, the addition of controls for marital status and education had much more of an effect on the estimated magnitude of $b$ than the addition of non-wage income alone did. This result indicates that as more constraints and characteristics other than unemployment are controlled for, the accuracy

<table>
<thead>
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<th></th>
<th>OLS</th>
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</tr>
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<tbody>
<tr>
<td>Age Controls</td>
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<td>0.090</td>
</tr>
<tr>
<td></td>
<td>[7.2%]</td>
<td>[7.1%]</td>
</tr>
<tr>
<td>Age and Non-Wage Income Controls</td>
<td>0.091</td>
<td>0.088</td>
</tr>
<tr>
<td></td>
<td>[7.8%]</td>
<td>[7.7%]</td>
</tr>
<tr>
<td>Age, Non-Wage, Observed Controls</td>
<td>0.074</td>
<td>0.071</td>
</tr>
<tr>
<td></td>
<td>[15.8%]</td>
<td>[15.2%]</td>
</tr>
</tbody>
</table>

N=351
Note: P-Values are in brackets.
The magnitude of $b$ estimated in Table 7 is very similar for both OLS and probit estimations of equation (12), with the probit estimates slightly lower than those of OLS. Overall, the estimates of $b$ range from .071 to .093, indicating that if one of a pair of otherwise similar sons had a father who was unemployed, this son would have a 7%-9% higher probability of being unemployed than he would otherwise. When only age and non-wage income are included in the equation, these results are statistically different from zero at the 10% level, but addition of observed controls for the son makes the P-Value equal to 15.2%, meaning one cannot say with confidence that this result is statistically different from zero.

Should we take these findings to mean that sons of unemployed fathers are more likely to be unemployed themselves? First it should be noted that many short spells of unemployment go unreported. This fact, in conjunction with the problem measurement error causes in the case discrete variables are used, could mean that the results reported here are severely biased. While it is difficult to know with certainty the direction of bias, it seems reasonable that this bias is downward. If this is the case, it seems fair to say that unemployment of sons is not independent of the unemployment of their fathers.

To further test whether the unemployment experiences of fathers and sons were independent, all conditional probabilities of whether a father was unemployed or not were estimated. Table 8 presents these results. The variable used in these calculations is the same as that in the regression equations, namely whether an individual was ever unemployed over a five-year period. The probabilities calculated were, conditional on the unemployment experience of a father, what is the probability of their sons being unemployed?

<table>
<thead>
<tr>
<th></th>
<th>father unemployed</th>
<th>father not</th>
</tr>
</thead>
<tbody>
<tr>
<td>son unemployed</td>
<td>0.32</td>
<td>0.23</td>
</tr>
<tr>
<td>son not</td>
<td>0.68</td>
<td>0.76</td>
</tr>
</tbody>
</table>

Note: Cells in table represent conditional probabilities.
Overall, the population average for sons is a 25% chance of being unemployed. However, conditioning on the unemployment experience of the father, those sons with unemployed fathers had a 32% chance of being unemployed as compared with a 23% chance if their father was not unemployed. Therefore, these results support the notion that having a father who was unemployed increases one’s own chance of becoming unemployed by 9%, as found in the estimation of equation (12). The mechanism for this intergenerational transmission is unclear, but it could result either from intergenerational constraints, or from preferences pertaining to unemployment that are cultivated within the family. Either way, this result surely suggests that, for sons ought to prefer having fathers who never are unemployed!

B. Mothers and Daughters

1. LABOR FORCE PARTICIPATION

Table 9 displays both OLS and probit estimates of \( b \) as specified by equation (11). This equation regresses the variable equal to one if a daughter worked for pay at any time during the period between 1989 and 1993 on the same variable for mothers between 1968-1972, using varying sets of controls. An important note is that the quantity reported for probit estimation of equation (11) is the derivative of the probability that a mother works. For comparability of OLS and probit results, the P-values of the estimates of \( b \) were reported. The first set of results, which only controls for age, is presented in the first row of Table 9. The second set of results includes controls for age as well as the educational level, marital status, and age of children of the daughter. OLS estimates of \( b \) range from .021 to .024, whereas the probit estimates range between .014 and .019. None of these estimates are statistically different from zero. The magnitude suggested by these results is smaller than might be expected, as it implies that probability of a daughter working increases by only 2.1% when their mother works. However, it should be noted that it is likely that the results estimated were downwardly biased, and as such this
estimate could represent a lower bound for the true probability of a daughter-working conditional on the work decisions of their mother.

In addition to analysis of the decision to work at any point in one’s five years of observation, this paper estimated equation (11) for a single year of data for both mothers and daughters using a probit functional form model. A large majority of these estimated coefficients were negative, with as great a magnitude as \(-.16\), SE=0.478. The overwhelming number of negative estimations of b prompted further analysis of the data concerning a daughter’s decision to work. What was discovered was that in a given year, almost 85% of daughters chose to work, compared with approximately 50% of their mothers. This finding is consistent with the trend of increasing female labor supply across generations, but is also more of an increase than most research indicates. Indeed, the finding that 93% of daughters worked for pay at some point during their five years of observation is unusual. When compared with the OECD’s (1999) estimate of 72.7% for the United State’s employment population ratio in 1991, a sample with 84% of daughters working in that year is clearly biased towards daughters who work. Possible explanations for why this sample had such a large percentage of working daughters are explored in part C of this section.

Table 10 presents conditional probabilities for whether a daughter works or not for both mothers who worked and those who did not. These probabilities are estimated as, given a mother’s decision to work positive hours during any of the years in the period of 1968-1972, what is the likelihood that their daughter chooses to work? As noted before, the overwhelming majority of daughters chose to work for pay, regardless of their mother’s decision. Despite this, the findings support one’s intuition the daughter’s are more likely to work if their mother chose to work. One can see this by comparing the

<table>
<thead>
<tr>
<th></th>
<th>OLS</th>
<th>Probit</th>
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<tbody>
<tr>
<td>Age Controls</td>
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<td>0.019</td>
</tr>
<tr>
<td></td>
<td>[47.1%]</td>
<td>[47.8%]</td>
</tr>
<tr>
<td>Age and Observed Controls</td>
<td>0.024</td>
<td>0.014</td>
</tr>
<tr>
<td></td>
<td>[38.8%]</td>
<td>[54.8%]</td>
</tr>
</tbody>
</table>

N=384

Note (1): P-values are reported in the brackets.
Note (2): Observed Controls include marital status, education, and the presence of children.
percentages in each row. For example, if one’s mother chose to work, one has a 6.2% probability of choosing not to work, compared with 8.1% if one’s mother did not work. However, these probabilities are based on a very small sample—only 26 total daughters did not work at any point between 1989-1993! In fact, of non-working daughters, more total daughters had mothers who did work (17 total versus 9). However, with only 26 observations, it is difficult to know how accurate the distribution of these observations was.

To further explore the results of Table 10, conditional probabilities of participating in the labor force for various numbers of years were estimated. These results are presented in Table 11. The variable in this table represents the total number of years a woman worked for pay over the five years observed. Since the variable is constrained to be between one and zero (one meaning an individual worked in all five years, zero indicating an individual never working for pay), each year worked adds 0.2 to the value of the variable for a given individual. Consistent with Table 10, within each column, the most likely result is that the daughter works in all five years. The interesting finding in the table comes from analyzing the second most likely outcome in any column. While for daughters whose mothers working a maximum of two of the five years the most likely outcome after working every year is to work in four of the years, for daughters whose mother worked all five years, the second most likely outcome is to not work in any of the five years! As in Table 10, a small sample size of daughters not working could have significant effects on the estimated probabilities, but these findings are not as one might expect.

2. Annual Hours of Work

Table 12 displays OLS and IV estimates of equations (8) and (9) for annual hours of work of mothers and daughters. Because almost 15% of daughters did not participate in the labor force in a given year, the sample of mothers and daughters used in these estimates was constrained to only those daughters working in all five years. This
restriction reduces the sample from 385 mother-daughter pairs to 277. This constraint was chosen to be able to observe more combinations of years of daughters and mothers, and to have consistency in the sample of mothers and daughters observed when doing so.

The results in the first five columns of Table 12 are estimated for a single year of annual hours of work of daughters, observed in \( t=1991 \), regressed on a single year of annual hours of mothers, where \( s=1968, 1969, \ldots, 1972 \). The column heading indicates the year of the annual hours of work of the mother. The first row only includes age controls, whereas the second adds controls for the marital status and education of the daughter, as well as whether she has young children. Using only age controls, the estimates of \( \hat{\alpha} \) range from -.039 to .02. None of these estimates are statistically different from zero. When analyzing estimates of all possible combinations of years for annual hours of work daughters and mothers, the estimates are very similar, ranging from -.058 (\( t=1993, s=1972 \)) to .038 (\( t=1992, s=1968 \)). Again, none of these estimates are statistically different from zero. The addition of controls for marital status, education, and children produced similar results. The estimates for \( t=1991 \) range from -.038 to .016 (as compared with -.039 to .02). Results in other years are similar.

Upon examining the results, the fact that both positive and negative estimates of \( \hat{\alpha} \) were reported is immediately surprising; the assumption would be that the relationship between annual hours of mothers and daughters would be positive. Which sign represents the true relationship between annual hours of work of mothers and daughters? First, it should be noted that none of these estimates were statistically different from zero, meaning that the sign of all these estimates cannot be regarded with much confidence. Second, it is important to recall that these estimates are subject to a substantial downward bias due to the measurement error and transitory movements in annual hours in a given year. Finally, all of the negative estimates of \( \hat{\alpha} \) were from estimations where \( s=1971 \) or
1972, meaning the mother was observed in 1971 or 1972. Is there a unique feature in the data for mothers in these years that could explain the negative correlations estimated?

One hypothesis might be that mothers worked uncharacteristically high or low hours in those years. Comparison of descriptive statistics of all five years of annual hours of work of mothers was revealing. While in 1968 and 1969 the mean annual hours of work for mothers was 638 and 695 hours respectively, in 1971 and 1972 these numbers were 763 and 732 hours. The standard deviation of annual hours of work was also higher in 1971 and 1972. This seems like a reasonable explanation, until one notes that in 1970 the mean annual hours were 749, although the standard deviation was slightly lower than in both 1971 and 1972. Examining differences in the maximum annual hours worked across years provides similar results. Overall, examination of the descriptive statistics of the annual hours of work of mothers seems to show that the higher mean and standard deviation of annual hours in 1971 and 1972 could be a partial explanation, but not the entire story.

Part of the explanation for the negative association between annual hours of work of mothers and daughters when mothers are observed in 1971 or 1972 could be that, in the early 1970s, the United States went through an economic recession. Such a recession could affect the work behavior of mothers in several ways. First, mothers already working could lose their jobs, or have their hours reduced. Conversely, if the spouses of married women lost their jobs, women might be forced either to begin working even if they had not been previously, or to work more hours. Based on the comparison of average annual hours of work of all mothers across years, it seems that women were

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<td>Age and Observed Controls</td>
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<td>0.016</td>
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</table>

N=278

Note (1): Standard error estimates are in parentheses.
Note (2): Personal controls include marital status, education and presence of children for the daughter.
Were a higher percentage of them working as well? Indeed, this seems to be the case, as between 53-58% of mothers worked in 1971-72, as compared with 46-52% in 1968-1969. The most important point is that all these statistics indicate that, in 1971 and 1972, women’s work behavior was changing to reflect the economic situation at the time. This means that in 1971 and 1972, women could be working hours very uncharacteristic for them, implying that their values of $v_{ips}$ (transitory fluctuation) for those years could be very high. The result of uncharacteristically high transitory fluctuation in annual hours of work in a given year would be a more downwardly biased estimate of $\tilde{n}$ than found in other years.

To address the issue of the errors-in-variables bias in the previous estimates of $\tilde{n}$, annual hours of work for mothers using more long-term measures were estimated, as outlined by equation (9). The results are presented in the sixth column of Table 12, using both age controls and age controls with the addition of observed characteristic controls for daughters. These results are smaller in magnitude than those using a single year of data for mothers and daughters, with estimates of -.005 and -.009. Standard errors were comparable to the estimates from single year estimates. In contrast to the case with father-son pairs, when using five-years of observations for mothers and daughters, the relationship between the annual hours of work of mothers and daughters is either obscured or ceases to exist. Considering that with a single-year of data, however, both positive and negative estimates of $\tilde{n}$ were found, this result is not particularly surprising. While the downward bias of the estimation of $\tilde{n}$ is expected to be smaller when using five years of data, the averaging of data across five years with different economic conditions may make the estimation of the equation using more long-term labor supply data more difficult.
As a final method to address the errors-in-variables bias, instrumental variables were applied to equation (8). These results are reported in the seventh column of Table 12. The instruments used were whether, in 1970, a mother had a child under the age of six or between the ages of six and eighteen. All IV estimates were for $t=1991$ and $s=1970$. Estimates of $\bar{n}$ range from .288 to .333. None of these estimates are statistically different from zero. They are, however, of a much higher magnitude than those estimated by OLS. These results indicate that the true association of annual hours of work between mothers and daughters may actually be positive, although the standard error in these estimates makes it difficult to say this with confidence.

To see if the lack of association found between annual hours of mothers and daughters could be explained by independently examining annual weeks worked and average weekly hours (as annual hours of work is the product of these two quantities), estimates of $\bar{n}$ for these variables were estimated. Table 13 presents these results. For all estimates, controls for the age of mothers and daughters, as well as marital status, education, and presence of children of various ages were included. Because weeks worked and average weekly hours were reported only within a range of weeks for each individual in 1968 and 1969, only results from 1970-1972 were estimated. Additionally for more long-run estimates of the weeks worked and average weekly hours, only three-year averages of 1970-1972 for mothers were used.

For weeks worked, the one-year estimates of $\bar{n}$ ranged from .027 to .03, none of which were statistically different from zero. When looking at estimates using three-year averages of labor supply data the estimates were similar, ranging from .003 and .017, estimates lower than with a single year of data. However, as discussed with annual hours of work, this result could be explained by the changing economic conditions in 1971 and 1972, meaning that the average were more likely to deviate from long-run measures of weeks worked than more short-run measures. Although no estimates were statistically different from zero, all estimates were positive in magnitude, implying a likely positive relationship of annual weeks worked between mothers and daughters.

The results for estimates using average weekly hours range from -.043 to .005. Again, none of these are statistically different from zero. It is interesting that negative relationships were measured, however, as this is the opposite of the estimates found for
fathers and sons. Whereas for fathers and sons, the correlation of average weekly hours was positive and that of weeks worked was essentially zero, the reverse was true for mothers and daughters. Three-year measures for average weekly hours were similar to those estimated with one year of data, ranging from -.025 to -.022.

The final column of Table 13 presents IV estimates of equation (8) with $t=1991$ and $s=1970$, as was done with annual hours of work. While neither estimate was statically different from zero, the magnitude of the $\bar{n}$ estimated was greater than that using OLS or long-run measures for these variables. For weeks worked, the estimate was .112 with SE=.325. For average weekly hours, the estimate was -.020 with SE=.427, a result almost identical to that found by OLS, with a higher SE. While it is difficult to draw many inferences from these results, as none can be reported as positive or negative with confidence, their similarity to estimates of $\bar{n}$ found by OLS indicates that these estimates are close to the true associations in the sample population.

To further evaluate the accuracy of these estimates, these results were compared with those found by Altonji and Dunn (1991). In their work, Altonji and Dunn reported estimates for annual hours between .063 and .128, although only the estimate of $\bar{n}=.091$ was statistically different from zero. These results fall in the middle of the range of estimates found in this work, which ranged from -.038 to .333. However, the high standard errors of the estimates both in this paper and the work by Altonji and Dunn indicates that further work needs to be done to truly evaluate the association between annual hours of work of mothers and daughters.

The comparison of the estimates of $\bar{n}$ reported by Altonji and Dunn for annual weeks worked provide striking differences. While the estimates in this paper range between .003 and .112, the results in Altonji and Dunn range between .108 and .522! Additionally, Altonji and Dunn’s results all were statistically positive. For average weekly hours, their estimates ranged from .036 to .137, as compared with -.043 to .005. As in this paper, none of the estimates of $\bar{n}$ for average weekly hours were statistically different from zero.

While many of the results found by Altonji and Dunn were similar to those found here, some important differences remain. What could explain these discrepancies? An important difference in the estimation done by Altonji and Dunn was their use of logs of
annual hours of work, weeks worked and, and average weekly hours. This estimation implicitly excludes mothers who did not work from the sample. However, in this paper, such mothers were included in the analysis. Could the exclusion of non-working mothers lead to different results?

To explore this hypothesis, all mothers who did work positive hours between 1968 and 1972 were excluded from the sample, and equation (8) was re-estimated. This restriction reduced the sample to only 92 observations. Using this new sample of working mothers and daughters, the estimates of $\hat{\eta}$ for annual hours of work range from .122 to .188, slightly higher than the estimates of Altonji and Dunn. Interestingly, many of these results were actually statistically different from zero and positive. This result implies that it is very probable that the association of annual hours of work of mothers and daughters is in fact positive. The changing composition of mothers working in each year as well as the extremely high percentage of daughters working in a given year could why such a result was not found in the original analysis of this paper. Regardless, this result is enlightening, and indicates that the association very well may be positive. In general, it seems more appropriate to include mothers who do work, as if one only includes working mothers and daughters in their sample, the question being examined changes to what the association of annual hours of work amongst working mothers and daughters when we would like to know the relationship amongst all mothers and daughters.

When analyzing the association of weeks worked amongst working mothers and daughters, the results are more similar to those found in Altonji and Dunn, although still smaller in magnitude. In this case, with the sample of 92 mother-daughter pairs, estimates of $\hat{\eta}$ range from .11 to .13, closer to the instrumental variables estimate found previously. Additionally, these estimates were almost statistically different from zero at the 10% level (the P-Values are approximately 11%). Analysis of average weekly hours with the reduced sample did not significantly alter any of the results found previously. Overall, these results seem to support the finding, that amongst working mothers and daughters, annual weeks worked is the greater source of variation in annual hours of work than average weekly hours, a result directly contrasting that found for fathers and sons.
C. Implications of Empirical Results

Given the notion that family socialization heavily impacts one’s tastes for work, the results presented in this section may seem somewhat surprising. While many of the intergenerational relationships in work behavior for fathers and sons did exhibit a positive association across generations, this result was not found for mothers and daughters amongst which no intergenerational associations in work behavior were found to be statistically different from zero. This result may be the most surprising, as intuition would tell us that female pairs might have stronger intergenerational links, as the variation in work behavior is much greater amongst females. What could explain the results found in this section? Should we take the results to mean that family socialization is not an important factor in the formation of tastes for work?

First, many of the intergenerational associations for father-son pairs were found to be statistically different from zero and positive, although many of a magnitude smaller than might have been assumed. The estimated correlation in annual hours of work was found to be approximately .09, and was found to be statistically positive at the 5% and 10% levels. This result indicates that sons of fathers who work more tend to do so as well. While the magnitude of the correlation is perhaps smaller than one might assume, the results imply that the sons of fathers who work 100 more annual hours, work 9 more annual hours themselves than they otherwise would based on their other observed characteristics. This may not seem like much, yet translated into a fifty-week work year, this result indicates that these sons would work ten extra minutes a week than they would otherwise.

Does this correlation hold for all levels of annual hours of work of fathers? To explore this question, probabilities of the annual hours of the son being in different quintiles of work hours amongst their peers conditional on the quintile of the annual hours of work of their father were estimated. The results are presented in Table 14. The bold cells represent the quintiles of the annual hours a son was most likely to work conditional on the quintile in which his father was. The interesting result in this analysis is that while for a given quintile of annual hours of the father there is a wide distribution
of sons across quintiles, with the exception of fathers in the fifth quintile, sons were most likely to be in the quintile above that of their fathers. Also interesting was that the highest probability in the entire table, by at least 6%, was that of sons and fathers both being in the highest quintile of work hours amongst their peers. In this category, sons whose father worked annual hours in the highest quintile had a 35% probability of also being in the highest quintile of work hours amongst their peers. This lends support to the notion that “hard-working” fathers often have sons who work a significant number of hours.

Upon examining the source of the positive correlation in annual hours between fathers and sons, the most likely explanation is in choices made in the average weekly hours worked. Correlations in annual weeks worked were plagued by a lack of variation and no positive correlation was found. On the other hand, statistically significant correlations in average weekly hours of the magnitude of .085 were found. This result indicates that a father working 10 more hours a week would have son who would work .8 more hours a week than he would otherwise. Translated into minutes, this means that a son worked 9.6 additional minutes per week worked. This result is almost identical to that found for annual hours of work, where a father working an additional 100 hours a year over fifty weeks, i.e. 2 more hours a week, would have a son who worked an approximately ten more minutes a week. As such, it seems plausible that the correlation in annual weeks worked could be minimal, as most of the correlation in annual hours of work can be accounted for by correlation in average weekly hours.

In contrast to the generally positive associations in labor supply behavior found for fathers-son pairs, few positive associations of statistical significance in work behavior for females were found. In magnitude, the increased probability of a daughter working if

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<th>Work quintile of Father</th>
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<th>2nd</th>
<th>3rd</th>
<th>4th</th>
<th>5th</th>
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<td>1st</td>
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<td>0.171</td>
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<td>0.214</td>
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</table>

N=348

Note: Values in cells represent conditional probabilities.
her mother worked was approximately 2%, but this result could not be reported to be
different from zero with confidence. In fact, analysis of all conditional probabilities of
whether a daughter will work based a mother’s labor force status shows a daughter’s
decision of whether to work to be independent of the decision of her mother.

The analysis of annual hours of work, annual weeks worked, and average weekly
hours was similar, with no associations found to be statistically different from zero and
some reported to be negative. However, when analyzing only mothers and daughters
working positive hours over all five years observed, the association between annual hours
and annual weeks worked of mothers and daughters are both positive and statistically
different from zero. While the sample size for this estimation was small (only 92 mother-
daughter pairs) it does indicate that the association in annual hours of work and weeks
worked is positive. Also, it seems that weeks worked explain much more of the variation
in annual hours of work of mothers and daughters, a result contrasting that found for
fathers and sons. Finally, it seems to suggest that part of the difficulty in estimating
correlations in the original sample with any confidence might derive from fluctuations in
the labor force participation of mothers.

Besides results found among working mothers and daughters, why were so few
positive associations found among this sample? Are we to assume that, in the general
population, work behavior of mothers has no association with that of their daughters?
This result seems unlikely. Instead, it seems that characteristics of the sample of
daughters in this research are the primary reason no association in the work behavior of
mothers and daughters was found. This is not to say that other research will not find the
association in work behavior between mothers and daughters is minimal, but based on the
sample used, one cannot in confidence say this is the true nature of work associations in
the population. The sample in this paper had an overrepresentation of daughters who
worked. In any given year, almost 85% of daughters worked, as compared with the
population rate of 70% at that time. Most likely, this overrepresentation of working
daughters in the sample obscured any relationship in the work behavior of mothers and
daughters, especially in the examination of the probability of labor force participation of
daughters. Even in the analysis of associations of annual hours amongst the sample of
daughters who worked, however, the overrepresentation is a problem. If a higher than
usual percentage of daughters are working positive hours while their mothers are not working at all, the true relationship between annual hours of work of mothers and daughters is obscured. As such, this problem of sampling most likely caused the inability of this analysis to find associations in work behavior between mothers and daughters.

What might have produced a sample with such a large percentage of working daughters? The best answer is that working daughters were more likely to respond in the original survey. The criteria for being included in the sample as a daughter were not especially constraining: one had to have lived in a household with a female in the sample in 1968, one had to be between the ages of 25 and 50 in 1989, one had to live in a household separate from their parents by 1989, and one had to report labor supply data in all five years of the survey. Additionally, only the eldest daughters meeting these criteria were retained in the sample. None of these requirements seems very restrictive, as 1800 daughters lived in the households of females in the sample in 1968, the age requirement is the same as for mothers, and by age 25, most individuals would have moved away from home anyways.

One possible explanation is that restricting the sample to the eldest daughters biased the sample towards daughters who work. Anecdotally, the eldest child is the most responsible, hard-working individual in the family. If for daughters, the difference in birth order made a large difference in one’s decision to work, this sample could over-represent working daughters. Further research could include all daughters meeting the sampling criteria to examine this possibility. However, it seems unlikely that the inclusion of younger daughters would cause the percentage of working daughters to decrease substantially, as no research seems to indicate that the work probabilities amongst older and younger siblings differ enough to bring the percentage of working daughters from 85% to 70%.

It is possible that the restriction of reporting five years of labor supply data caused a bias towards working daughters. However, there is no obvious intuitive reason why an individual responding in 1989 and not 1990 would be less likely to work. Additionally, a large fraction of the daughters already had not responded by 1989, as the overall non-response rate in the survey by 1989 was almost 45% (Hill 1991). The more likely
concern is the non-response rate of daughters as a whole by 1989, as annual attrition rates are very low, yet added over time are more of a concern.

One theory of the higher response rate of working daughters could be related to marriage. Unmarried women are more likely to work, so an under representation of married daughters in the sample could explain the overrepresentation of working daughters. Only 70% of daughters were married, as compared with almost 90% of mothers. A large portion of the decrease in marriage across generations is surely explained by the changing demographics, but it seems reasonable that married women might be less likely to respond. Many women change their names once they get married, and as such, they might be more difficult to track over time. However, marriage rates of daughters across the five years observed were relatively stable, making it unclear that married daughters are more likely to respond. Overall, further research is needed to determine why working daughters are over represented in this sample, but regardless of the cause, in this research, it is clear that this bias obscures any associations between the work behavior of mothers and daughters.

V. CONCLUSION

In the analysis of labor supply, researchers consistently find a wide variation in the annual hours of work by individuals. Only a strikingly small percentage of this variation can be explained by observed characteristics of individuals and their work experiences, is it average hourly wage, age, marital status, race, education, etc. This finding has prompted the conjecture that preferences play an important role in the labor supply decisions made by individuals. The question has been, how are these preferences formed, and how much of the variation in hours across individuals do they explain? A major hypothesis is that family socialization is an important source of the formation of preferences for work. This research takes first steps to analyze this hypothesis, using intergenerational data from the Panel Study of Income Dynamics to examine intergenerational associations in work behavior. Overall, positive associations were found in the work behavior of fathers and sons, though of a smaller magnitude than might have been assumed. For father-son pairs, the intergenerational correlation in annual hours of work was .09, almost all attributable to correlations in average weekly hours.
Additionally, sons of fathers who were unemployed were 9% more likely to be unemployed than they would be otherwise. While few statistically positive associations were found amongst females, this result is most likely caused by the overrepresentation of females in the sample of this paper. Further research must be done to better evaluate the intergenerational associations of work behavior of mothers and daughters.

An obvious limitation of this study is its reliance on one data set and its small sample size. Similar research done by Altonji and Dunn (1991) using a dataset from the National Longitudinal Surveys that had a larger sample size produced similar results to those presented here, however. Therefore, to more accurately estimate intergenerational correlations in work behavior, data over a longer time period is needed, as with more permanent labor supply data, associations between generations can be better estimated. Additionally, models accounting for differences in the group of individuals staying in the survey need to be included, especially in the case of females.

As this research is only a first step, further research along several dimensions examined in this paper could be done. This most obvious is to find sampling criteria that ensure that the percentage of working daughters in the sample better mirrors that of the general population. Section IVC gives suggestions of how one might best pursue this analysis. Another idea for further research is to not only study the associations of father-son pairs and mother-daughter pairs, but of all pairs of parents and children. While one would imagine that children most closely mimic the behavior of their parent of the same gender, this does not mean other associations amongst parents and children would not exist, and as such, this analysis could provide further insight into socialization within the family.

As a final note, it is important to realize that the results of this analysis are only relevant to the time and generations for which they were measured. With the ever-changing demographics in our nation, it is entirely possible that if this analysis were carried out for later cohorts, the results would be very different. If the hypothesis that positive associations in work behavior between generations are largely a result of family socialization were true, the changing nature of the American family would most certainly cause the intergenerational associations in work behavior to change. Differences in the labor force participation rates of women, rates of marriage and divorce, number of
children in families, as well as the changing racial make-up of our society leave it a very
different nation than it was in 1968. It is a nation with changing values and priorities.
These values are at the core of family socialization. As our nation and families changes,
one can expect that the associations of work behavior of parents and their children will
change as well.
### APPENDIX

#### Appendix Table One

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<th>Year of son's annual hours of work</th>
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Note: all equations use age + budget controls
N=348

#### Appendix Table 2
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N=350
Note: all equations use age + budget controls

REFERENCES


