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The Covariance Structure of Intrafamily Earnings, Rising Inequality and Family Labor Supply

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Abstract

This paper uses an intertemporal family labor supply framework to study the evolution of individual and family earnings during the period of rapidly rising wage inequality in the early 1980s. Working couples have strongly positively correlated labor market outcomes. However, the correlation of a wife's wages or earnings with her husband's outcomes in earlier or later years is roughly constant, implying that the inter-family correlation is entirely attributable to permanent factors. This finding motivates a very simple intertemporal family labor supply model in which husband's and wive's labor supply decisions are linked solely through the family budget constraint and permanent taste factors. Estimation results for this model yield a small estimated intertemporal labor supply elasticity for married men (0.05), but a larger elasticity for married women (0.40). The magnitudes of these estimates imply that behavioral labor supply responses explain relatively little of the rise in annual earnings inequality for married men, but over 50 percent of the modest rise in inequality for married women, and 20 percent of the rise in family earnings inequality in the early 1980s.

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Introduction

There is now an extensive literature documenting the rise in the cross-sectional inequality of labor market outcomes in the United States during the 1980s.[1] Most of this literature has focused on the dispersion in rates of pay across individual workers, and on the contributions of such factors as trade and technology to the growth in the inequality of wages within and between various demographic groups.[2] Despite this focus, much of the policy interest in labor market inequality concerns its effect on the inequality of family well-being. While it is known that family income inequality has risen as much or even more than individual wage inequality, there is a surprising absence of research on the mechanisms underlying this link.[3]

There are three key labor market features that affect the link between the dispersion in rates of pay and the inequality of family well-being. The first is the extent to which wage inequality is attributable to permanent versus transitory shocks.[4] A rise in the variance in transitory income shocks will tend to have far less serious welfare consequences than a comparable rise in permanent inequality. A second factor is the degree to which wage changes of workers in the same family are strongly or weakly correlated. Any positive correlation between husbands' and wives' wage shocks will compound any rise in individual wage inequality at the family level. A third issue is whether families change their

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[1] See Levy and Murnane (1992) for a comprehensive survey. Prominent studies include Cutler and Katz (1991), Gottschalk and Moffitt (1994), Juhn (1992), Juhn, Murphy and Pierce (1993), Katz and Murphy (1992), Murphy and Welch (1992), and Pierce and Welch (1992).

[2] See Katz and Autor (1999) for a recent survey. Although most studies implicitly model the rate of pay, they frequently use annual or weekly earnings to measure inequality.

[3] For example, see Cutler and Katz (1991), Hyslop (1995), Juhn and Murphy (1996), and Pierce and Welch (1992).

[4] Most analyses have concentrated on cross-sectional inequality. One exception is Gottschalk and Moffitt (1994) who use panel data to analyze the persistence in cross section differences in log annual male earnings in the two periods 1970-78 and 1979-87. Using simple error components specifications, they find that the total variance of earnings increased 50 percent between 1970-78 and 1979-87, and the permanent variance accounted for about two-thirds of the total in each period, implying that the permanent and transitory variances increased proportionately.

hours in response to shifts in the hourly rates of pay of individual family members. In the absence of behavioral responses, hourly wage inequality will be transmitted mechanically into individual annual earnings inequality, and family earnings inequality will evolve according to the changes in the husbands' and wives' earnings and the degree of assortiveness of husband-wife matching. Labor supply responses can work in several directions to accentuate or mitigate these mechanical effects. For example, intertemporal substitution effects in which each spouse's hours respond positively to changes in their wage rates would tend to accentuate increases in inequality, while income effects might mitigate the increase. Also, depending on the nature of household preferences, a rise in one spouse's wage could also have an indirect effect on the other's labor supply, and may raise or lower inequality.

This paper uses an intertemporal labor supply framework to examine these three issues and study the link between rising wage inequality and family earnings inequality for two-earner families in the period from 1979 to 1985. The analysis focuses on this relatively short period for three reasons. First, the most dramatic recent changes in labor market outcomes appear to have been concentrated in this period.[5] It is therefore important to understand how these changes affected families. Second, the magnitude of the increase in wage dispersion during the early 1980s suggests that it provides an ideal testing ground for identifying the labor supply responses of individuals and their spouses to changes in the financial incentives for work. Third, although the models used in the paper can be easily applied to a longer sample period, the data requirements for the model impose more stringent sample selection requirements the longer the period of analysis.

The paper begins by using longitudinal data from the Panel Study of Income Dynamics (PSID) to conduct a descriptive analysis of the variance-covariance structures of spouses' wages and their earnings, using a simple error components model of wages which allows growth in inequality over time.

[5] For example, Blank and Card (1993), and Juhn, Murphy and Pierce (1993) document that much of the change in wage dispersion during the 1980s was concentrated in the period 1979-85.

I first examine the autocovariances of individuals' wages, and focus on what fraction of the variance of wages is attributable to permanent versus transitory shocks. I then examine the cross-covariances of husbands' and wives' wages, and the degree to which these are attributable to permanent or transitory factors. Third, I examine the covariance structure of spouses' earnings and ask whether it is reasonable to assume that individuals' hours worked are independent of their wages. The important findings from this analysis are, first, that about one-half of the variances of wages are permanent; second, that the correlation of a wife's wages or earnings with her husband's outcomes in earlier or later years is positive and roughly constant; and third, the cross-correlation between spouses' wages is greater than between their earnings. The implication of the latter finding is that hours are not independent of wages, and the difference in the cross-correlations in wages and earnings are entirely attributable to permanent factors. Importantly, there is no intertemporal cross-substitution effect linking changes in one spouse's wages to changes in the other's hours.

I then present a simple intertemporal family labor supply model that is consistent with these conclusions, and combine this with an error-components model for the evolution of husbands' and wives' hourly wages. The model is fit to means, variances, and covariances of husbands' and wives' wages and earnings for continuously-employed couples in the PSID over the 1979-85 period. The results from the wage model imply that the rise in the cross-sectional inequality in husbands' hourly wages was roughly proportionately due to permanent and transitory factors and the permanent share was about one-half over the early 1980s, while for wives' the rise was largely due to transitory factors and the permanent share declined from 65 percent to 55 percent. Parameter estimates from the labor supply model imply that the intertemporal labor supply elasticity of husbands with respect to their own wages is close to zero, while for wives it is about 0.4 for wives. The estimates appear to be robust to alternative methods of adjusting for the non-representative nature of the sample of wives who are continuously employed over the 7-year sample period. The magnitudes of these estimated intertemporal elasticities imply that behavioral labor supply responses explain very little of the rise in annual earnings inequality for married men, but do explain over 50 percent of the modest rise in inequality for married women, and 20 percent of the rise in family earnings inequality in the early 1980s.

I: Individual and Family Earnings Inequality

In this section I first provide a brief overview of the trends in cross-sectional income inequality of individuals and families over the early 1980s. I then present a simple statistical model that can be used to quantify the key dynamic features of individual wage and earnings outcomes that affect the link between wage inequality and family well-being. This emphasizes the critical distinction between permanent and transitory wage components, and provides a natural framework for assessing the role of permanent factors in a non-stationary economic environment. I then use this statistical model to conduct a descriptive analysis of the variance-covariance structure of wages and earnings of husbands and wives in the PSID. The findings from this analysis lead to the specification of a particular intertemporal family labor supply model in section II, that is used to formally model the linkage between individual wage inequality and family earnings inequality.

Trends in US Cross-Sectional Earnings Inequality

As a backdrop for the analysis in the remainder of the paper, Table 1 summarizes the trends in cross-sectional inequality in individual and family earnings in the US economy during the first half of the 1980s.[6] Columns (1) and (2) summarize cross-sectional samples from the Current Population Survey (CPS), while columns (3) -- (5) summarize longitudinal samples from the PSID.[7] Throughout the paper

[6] Cutler and Katz (1991) estimate that 90 percent of total family income is labor earnings. Therefore, as a practical matter, an analysis of earnings inequality should provide an accurate account of income inequality.

[7] The appendix describes the sample selection criteria, and appendix table A1 summarizes some of the demographic characteristics of the samples.

I use the squared coefficient of variation of earnings to measure inequality.[8]

The results in column (1) pertain to samples of married couples with both spouses aged 18-60 in each year. Mean real earnings of men were approximately constant over the period 1979 to 1985, while earnings inequality increased substantially (38 percent). Most of these changes were concentrated in the 1979-82 period: mean earnings fell 8 percent and inequality increased 33 percent in this period. In contrast to the trends for married men, the earnings of married women rose 13 percent in the early 1980s, while inequality increased only slightly.[9] The trends for total family earnings are similar to those for men, with a modest rise in the level of earnings and a pronounced increase in inequality. The results in column (2), which hold constant the cohort composition of the sample, are essentially the same as in column (1).

The analysis in this paper uses longitudinal data for 1979-85 from the 1986 release of the PSID. I include in the sample both the random Census-based component of the PSID and the poverty-oversample. The PSID sample of husbands and wives described in column (3) pertains to continuously married couples who were aged 18-60 in 1980. In principle, this sample is comparable to the CPS sample in column (2), although it over-represents poor and nonwhite households. Moreover, the requirement that the PSID couples be continuously married introduces some additional selectivity relative to the CPI samples which only require marriage at a point in time. The trends in the inequality of earnings in the PSID sample are similar to the patterns observed in the CPS samples, however the <u>level</u> of inequality is lower. In particular, the squared coefficients of variation of husbands' and wives'

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[8] The variance of log earnings has perhaps been the most commonly used measure of dispersion in the recent literature -- e.g., Gottschalk and Moffitt (1994) and Juhn, Murphy and Pierce (1993). The squared coefficient of variation provides a first-order approximation to this measure, and is used in this paper because of the interest in analyzing spouses' individual and joint earnings inequality. Levy and Murnane (1992) discuss the properties of alternative measures of earnings inequality.

[9] One reason for the smaller increase in inequality in female earnings in this period is that mean annual hours worked by females increased during the period without any increase in the variance of hours worked, thus counter-balancing the increasing variance of female wages.

earnings rose by 46 percent and 10 percent respectively. As in the CPS, the rise in male inequality was concentrated in the 1979-82 period. Finally, family earnings inequality rose by about as much as for husbands.

In column (4), the PSID sample is restricted to couples who report positive male earnings in each year of the sample period; in column (5), the sample is further restricted to couples who report positive male and female earnings in each year. The trends in male earnings inequality are similar across the three PSID samples, although average earnings and inequality are slightly lower in the continuously-working couples sample. In contrast, average female and family earnings are significantly higher, and inequality is lower, in the third sample compared to the first two samples. However, the increase in family earnings inequality (24 percent) remains substantial in this sample.

The data in table 1 confirm that the substantial increases in male and family cross-sectional earnings inequality during the early 1980s were widespread across various samples. Comparable increases are observed using both cross-sectional and longitudinal samples, and using various selection criteria based on demographic and labor market characteristics of the samples, although selection on positive female earnings may affect the analysis of female and family earnings. I now turn attention to the persistence in wage and earnings differences observed in the cross-section, focusing on the PSID sample of continuously-working couples.

A Statistical Analysis of Persistence in Cross-sectional Inequality

It has long been recognized that persistent differences in individual or family income are more important determinants of the distribution of wellbeing than transitory shocks (e.g., Friedman, 1957, Modigliani and Brumberg, 1954).[10] This observation suggests that it is important to know whether the

[10] This literature has lead to an alternative approach to measuring inequality using the distribution of consumption. Cutler and Katz (1991) conclude that, although the level of consumption inequality is lower than income inequality, the trends over the 1980s are similar.

rises in annual earnings inequality in the early 1980s described in table 1 represent rises in permanent or transitory differences. To examine these issues, I now introduce a framework that focuses on the roles of the persistence in individuals wages, the relationships between spouses' wages, and the hours of work of spouses' in evaluating the sources of and persistence in earnings inequality.

In this context, the natural starting point is a characterization of the persistence in individual wage differences, and the simplest model for this purpose is:[11]

$$w_{it} = \Gamma_t + \alpha_i + u_{it} \qquad (1)$$

where w_{it} is the wage of individual i in year t; Γ_t is the mean wage across individuals in year t, which captures common aggregate effects; α_i is an individual-specific permanent component; and u_{it} is a transitory component. The wage components α_i and u_{it} are assumed to be orthogonal with $E(\alpha_i)=E(u_{it})=0$. The usual interpretation of α_i is that it represents the human capital or skill characteristics of the individual, such as their education and experience levels; while u_{it} captures the effects of transitory wage or labor market shocks.[12]

Equation (1) implies that cross-sectional wage inequality is $CV(w_{it})^2 = var(w_{it})/\Gamma_t^2 = \{var(\alpha_i)+var(u_{it})\}/\Gamma_t^2$. The immediate problem with specification (1) for studying inequality in a period such as the early 1980s is that, except for Γ_t, the only source of non-stationarity in wages is via the transitory component, u_{it}. Thus, an application of this model to the early 1980s would imply that the rise in cross-sectional inequality was due entirely to transitory factors. However, there is a wide body of evidence

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[11] This model is closely related to literature on longitudinal earnings processes for individuals. For example, see Hause (1980), Lillard and Weiss (1979), Lillard and Willis (1978), and MaCurdy (1982). Smith (1979) estimates error components models separately for male and family earnings to analyze the effect of female earnings on family inequality, and Card (1994) estimates an error components model for male wages. One important distinction with this literature is that, as the analysis here is also interested in combined spousal earnings, the specification is in terms of wage levels rather than logarithms.

[12] This model assumes there is a single index of skill -- i.e., the returns to different skill dimensions (e.g. education, experience, etc.) change proportionately over time. Also, note that the transitory component u_{it} may be correlated over time, however $cov(u_{it},u_{it+k}) \rightarrow 0$ as $|k| \rightarrow \infty$.

which suggests that at least some of the recent increase in wage inequality is attributable to increases in the importance of permanent factors, such as education.[13] A simple modification to equation (1) that allows non-stationarity in wages is:

(2) $$w_{it} = \Gamma_t + \theta_t(\alpha_i + u_{it})$$

where θ_t is a year-specific "factor loading" on the permanent and transitory components, and I now assume $var(u_{it})$ is constant over time. In this specification, the permanent and transitory components of variance expand or contract proportionally over time, according to changes in the return to skill, represented by θ_t. Different factor loadings for the permanent and transitory components can be easily introduced, and I pursue this formally in the next section.

Equation (2) provides the following three relevant predictions. First, cross-sectional wage inequality is $CV(w_{it})^2 = var(w_{it})/\Gamma_t^2 = \theta_t^2\{var(\alpha_i)+var(u_{it})\}/\Gamma_t^2$, which rises if θ_t/Γ_t rises. Second, permanent wage inequality is $CV^P(w_{it})^2 = \theta_t^2 var(\alpha_i)/\Gamma_t^2$, which is a fixed fraction of total cross-sectional inequality. Third, although the loading factors play an important role in both cross-sectional and permanent inequality, they do not affect the autocorrelations in individual wages in this specification. Specifically, the fraction of cross-sectional inequality which persists after k years, $corr(w_{it},w_{it+k})$, does not depend on θ_t. Also, noting that $cov(u_{it},u_{it+k}) \rightarrow 0$ as $|k|\rightarrow\infty$ implies that $corr(w_{it},w_{it+k}) \rightarrow var(\alpha_i)/\{var(\alpha_i)+var(u_{it})\}$, the permanent fraction of cross-sectional inequality. Thus, the autocorrelations in individuals' wages contain important information for understanding the extent of persistence in inequality.

Equation (2) can be simply extended to consider husbands and wives' bivariate wage process:

(3a) $$w_{mit} = \Gamma_{mt} + \theta_{mt}(\alpha_{mi} + u_{mit})$$

(3b) $$w_{fit} = \Gamma_{ft} + \theta_{ft}(\alpha_{fi} + u_{fit})$$

[13] For example, Juhn, Murphy and Pierce (1993) conclude that about 50 percent of the increase in inequality over the 1980s can be attributed to increasing returns to observable skill factors.

where the subscripts m and f are used to denote the husband and wife respectively, and i now denotes the family. A natural way to incorporate interactions between spouses' wages in equations (3a) and (3b) is to allow their permanent components to be correlated to characterize the "assortive" matching of husbands and wives, and also to allow their transitory wage shocks to be correlated to characterize the nature of transitory labor market shocks within families.

Equations (3a) and (3b) lead to the following implications for family inequality. First, although the loading factors affect the cross-covariances, they do not affect the cross-correlations between spouses wages. Second, assuming there is both positive assortive mating and positive correlation between spouses' transitory shocks, $corr(w_{mit}, w_{fit+k})$ is positive and declines as $|k|$ increases.[14] In fact, if spouses' transitory wage shocks are uncorrelated, then the cross-correlations between their wages will be constant over time and at different leads and lags, with a magnitude that reflects the degree of assortiveness in the earnings capacity of couples. Similar implications will hold if the permanent and transitory components have distinct loading factors, so long as these factors do not vary too differently.

Assuming that the wages of husbands and wives evolve over time according to equations (3a) and (3b), I now consider the implications for the evolution of earnings. This depends critically on how individuals respond to the permanent and transitory components of their own and their spouses' wages. As a base case, consider what might be called the "no labor supply" assumption that husbands and wives work fixed numbers of hours each year (γ_m and γ_f respectively), with no response to wages:

(4a) $$y_{mit} = \gamma_m w_{mit} = \gamma_m \Gamma_{mt} + \gamma_m \theta_{mt}(\alpha_{mi} + u_{mit})$$

(4b) $$y_{fit} = \gamma_f w_{fit} = \gamma_f \Gamma_{ft} + \gamma_f \theta_{ft}(\alpha_{fi} + u_{fit}).$$

In this case, spouses' annual earnings are simply scaled versions of their hourly wages, so that the correlation structure of spouses' earnings will match exactly that of their wages described above, and

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[14] That is, $corr(w_{mit}, w_{fit+k}) = \{cov(\alpha_{mi}, \alpha_{fi}) + cov(u_{mit}, u_{fit+k})\}/\{(var(\alpha_{mi}) + var(u_{mit}))(var(\alpha_{fi}) + var(u_{fit+k}))\}^{1/2}$. As $|k| \rightarrow \infty$, $cov(u_{mit}, u_{fit+k}) \rightarrow 0$, and $corr(w_{mit}, w_{fit+k})$ converges to the assortive mating component of correlation.

discussions of individual inequality could equally be conducted either in terms of wages or earnings. Moreover, equations (4a) and (4b) implies that permanent family earnings inequality will rise with the degree of assortive mating by couples, and that correlated transitory shocks within families will also affect cross-sectional inequality.[15]

In general, if individuals hours of work vary systematically with either their own or their spouses' wages, the predictions for earnings are ambiguous. For example, if individuals respond positively to changes in their own wages, earnings inequality will tend to increase relative to wage inequality. In contrast, the presence of income effects in hours decisions will likely mitigate the transmission of wage inequality into earnings inequality. I return to this issue in section II, after first discussing the characteristics of the empirical covariance structures of spouses' wages and earnings.

The Covariance Structures of Spouses' Wages and Earnings

I now turn to the empirical covariance structures of wages and of earnings for the PSID continuously-employed couples sample described in column (5) of table 1, to evaluate the predictions associated with equations (3a & b) for spouses' wages, and equations (4a & b) for their earnings in the absence of labor supply considerations. The full covariance matrices of wages and earnings are presented in appendix tables A2 and A3 respectively, while table 2 summarizes the structure of wages in panel A, and of earnings in panel B.

I begin with the covariance matrix for wages. First, the autocorrelation patterns of males' and females' own wages are quite similar. These show that about three-quarters of cross-sectional wage differences persist after 1 year, and more than 60 percent remain after 6 years. This implies that

[15] Total spousal earnings is the sum of (4a) and (4b). Thus, cross-sectional family earnings inequality is $CV(y_{mit}+y_{fit})^2 = \{(\gamma_m\theta_{mt})^2(var(\alpha_{mi})+var(u_{mit})) + (\gamma_f\theta_{ft})^2(var(\alpha_{fi})+var(u_{fit})) + \gamma_m\gamma_f\theta_{mt}\theta_{ft}(cov(\alpha_{mi},\alpha_{fi})+cov(u_{mit},u_{fit}))\}/(\gamma_m\Gamma_{mt}+\gamma_f\Gamma_{ft})^2$; and permanent inequality is $CV^P(y_{mit}+y_{fit})^2=\{(\gamma_m\theta_{mt})^2var(\alpha_{mi})+(\gamma_f\theta_{ft})^2var(\alpha_{fi})+\gamma_m\gamma_f\theta_{mt}\theta_{ft}cov(\alpha_{mi},\alpha_{fi})\}/(\gamma_m\Gamma_{mt}+\gamma_f\Gamma_{ft})^2$.

transitory wages are serially correlated and that, perhaps, about one-half of the cross-sectional inequality is due to permanent factors. Second, there is no systematic time-variation in the autocorrelations in individuals' wages at given lags, which implies that the relative persistence in wage differences is roughly constant over the period. This is consistent with the prediction from equations (3a) and (3b) that the increase in cross-sectional wage inequality over time is due to proportional increases in permanent and transitory factors.

The cross-covariances between spouses' wages, shown in table A2, are always positive and significantly different from zero. However, in contrast to the serial correlation in own-wages which decline with the lag length k, the cross-correlation between male and female wages are roughly constant at different leads and lags, and also across time (equal to 0.32 on average). This implies there is strong positive assortive matching by couples, and negligible correlation between spouses' transitory labor market shocks. For example, if one-half of the cross-sectional inequality in individual wages is due to permanent factors, and the cross-correlation between spouses' wages is equal to 0.3, then the correlation between their permanent wages will be equal to 0.6.

I now turn to the covariance structure of spouses' earnings. The autocorrelations in male and female own-earnings are again similar, although differences are more persistent in female earnings than in male earnings. However, the autocorrelations are generally <u>higher</u> in earnings than in wages, in contrast to the prediction from equations (4a) and (4b) that these will be the same. The average difference in correlations is about 0.12 for females and 0.03 for males, and these differences do not appear to vary systematically across time or lag. On the face of it, this suggests that both men's and women's hours respond positively to their wages in this sample, and these responses dominate any mitigating income effects that may exist and act increase earnings inequality at the individual level. In addition, the autocorrelations in female earnings tend to increase over time (see table A3) which, given the absence of a similar pattern in wages, suggests that persistence in hours worked by females increased over the period.

The cross-correlations between spouses' earnings are positive and roughly constant (about 0.25) across both time and lag differences, reflecting the effects of permanent rather than transitory factors.[16] However, the cross-correlations are lower in earnings than wages by about 0.07 on average. Thus, in contrast to differences in individual wage and earnings correlations, this implies the presence of an income effect in hours tends to dominate intra-family behavior and acts to mitigate the assortive mating effects of couples. Although the contrasting implications from the autocorrelations in individuals' wages and earnings and the cross-correlations between spouses' wages and earnings are not necessarily inconsistent, they are perhaps counterintuitive. One explanation which may reconcile these differences is the confounding effects of measurement error. For example, as measured wages are equal to reported annual earnings divided by reported annual hours worked, it is likely that there is greater measurement error in wages than earnings. Consequently, there will be greater attenuation bias in the correlations in wages than in earnings, and may explain the greater autocorrelations in individuals earnings than wages. For this reason, the smaller cross-correlations in spouses' earnings than in their wages is more robust than the reverse finding in individuals' own wage and earnings autocorrelations.

Finally, the autocorrelations in family earnings described in table 2 are higher than the corresponding correlations in individual earnings. This implies that the combined effects of positive assortive mating of couples and positive labor supply elasticities with respect to wages tend dominate the mitigating income effect on family earnings inequality, although this is subject to the same caveat concerning measurement error.

The descriptive analysis presented in this section has highlighted several important stylized facts for the modelling the link between dispersion in individuals wages and family earnings inequality. These

[16] However, there does appear to be a small but systematic increase in the cross-correlations between husbands' earnings in year t and their wives' earnings in t+k as k increases, which is consistent with compensatory behavior by women in response to their husbands' transitory earnings shocks.

facts can be summarized as follows. First, a substantial fraction of cross-sectional wage dispersion is transitory in nature, and it appears that this fraction was approximately constant over the period of interest. Second, spouses' wages are positively correlated, and the correlation appears to be substantially due to strong positive assortive matching by couples. Third, comparing the correlation structures of spouses' wages and their earnings reveals differences that imply families hours of work decisions play an important role in the link between wage and earnings inequality.

There are also several caveats to the analysis presented here. Most importantly, the analysis has been conducted in the absence of a model of labor supply decisions. That is, the analysis assumed that families hours of work do not respond to their wages, which appears to be an inadequate description of the data. In addition, the bivariate wage process assumed the growth in permanent and transitory inequality to be equal in order to obtain some predictions from the model. Although this appears to be a reasonable restriction on the evolution of individual wages, it is straightforward to relax this assumption and allow divergent changes in permanent and transitory inequality. Third, the analysis has ignored possibly confounding effects of measurement error in observed wages and earnings. Finally, the analysis uses a sample of continuously married couples, who each work throughout the period; this issue raises the question of how representative the results are. In the next section I formalize the framework introduced in this section, and extend it to address each of these caveats.

II: Intrafamily Earnings and Family Labor Supply

In this section I develop an econometric framework to analyze the link between wage and earnings inequality which extends the descriptive model in the previous section. I first formalize the model for spouses' bivariate wage process. Next, I develop a simple intertemporal family labor supply model, which is consistent with the stylized facts discussed in section I, to model the hours of work decisions of families. Finally, I examine the effect of non-random sample selection, and correct for the nonrepresentativeness of the sample associated with the labor force participation decisions of wives. The results from this analysis are broadly in line with the findings in section I, and help quantify the magnitude of the effects of interest.

An Empirical Specification for Wages

I begin by extending the bivariate process for spouses' wages in equations (3a) and (3b) to specify a statistical process to account for serial correlation in transitory wages, allow the non-stationarity in permanent and transitory wage components to vary, and to allow for random measurement error in observed wages. In particular, the empirical specification for males and females observed wages is

$$\text{(5a)} \qquad w^*_{mit} = w_{mit} + \eta_{mit} = \Gamma_{mt} + \theta_{mt}\alpha_{mi} + \phi_{mt}u_{mit} + \eta_{mit}$$

$$\text{(5b)} \qquad w^*_{fit} = w_{fit} + \eta_{fit} = \Gamma_{ft} + \theta_{ft}\alpha_{fi} + \phi_{ft}u_{fit} + \eta_{fit}$$

$$\rho_\alpha = \text{corr}(\alpha_{mi}, \alpha_{fi});\ u_{jit} = \rho_j u_{jit-1} + \upsilon_{jit},\ (j=m,f);\ \text{and}\ \rho_\upsilon = \text{corr}(\upsilon_{mit}, \upsilon_{fit}).$$

That is, the transitory wage components, u_{mit} and u_{fit}, are assumed to follow stationary first-order autoregressive (AR(1)) processes, with innovations υ_{mit} and υ_{fit}, and the correlation between spouses' transitory shocks is formalized by allowing the innovations to be correlated. Unrestricted time-varying factor loadings on each of the permanent (θ_{mt} and θ_{ft}) and transitory (ϕ_{mt} and ϕ_{ft}) components provides a flexible specification of non-stationarity in wages. Subsequently, I will refer to $\theta_{jt}\alpha_{ji}$ and $\phi_{jt}u_{jit}$ as the permanent and transitory components of wages respectively. Finally, I assume η_{mit} and η_{fit} represent measurement error in wages, which is uncorrelated across time and within families.[17] For identification purposes, α_{ji}, u_{jit}, η_{jit} (j=m,f) are assumed to be mutually orthogonal, with $E(\alpha_{ji})=E(u_{jit})=E(\eta_{jit})=0$. In

[17] Although I will interpret η_{mit} and η_{fit} as measurement error, these components will also absorb any purely transitory components of true wages; conversely, u_{jit} may also include serially correlated measurement error in spouse j's wages. Also, restricting the purely transitory errors to have constant variances implicitly assumes that measurement error in wages is absolute. Pischke (1995) finds that measurement error in earnings may be adequately modelled either as absolute or relative error.

addition, the initial period scale factors, θ_{j1} and ϕ_{j1} (j=m,f), are normalized to unity.

To evaluate the adequacy of the bivariate wage model (5a) and (5b), it is fit to the empirical covariance structure of the sample of continuously-working spouses' wages described in table 2. The model is estimated using generalized method of moments (GMM) techniques, and minimizes a weighted sum of squared differences between the theoretical covariance matrix of wages implied by the model and the sample covariance matrix.[18] Due to small sample bias associated with optimally weighted minimum distance (OMD) estimation of covariances structure models, I use a diagonal weight matrix with diagonal elements equal to the inverse of the sampling variances of the moments of the data.[19]

Table 3 summarizes the results from this exercise, which are broadly similar to the descriptive results discussed in section I. The GMM goodness-of-fit statistic is 88.95 (p-value=0.07) which implies the specification provides a reasonable statistical description of the wage data for this sample.[20] The variance estimates imply that, in 1979, about 50 percent of the variance of observed male wages is due to permanent factors, 30 percent is due to transitory factors, and 20 percent is measurement error. The corresponding estimates for females are two-thirds, 9 percent, and 24 percent respectively. Also, the correlation parameters associated with the AR(1) transitory wage components are over 0.8 for males and females; thus, there is substantial persistence in the transitory wages.

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[18] See Abowd and Card (1989), Chamberlain (1984), and Newey (1985) for discussions of these estimators.

[19] See Altonji and Segal (1996) for a detailed analysis of the problem, which arises from positive correlation between the estimates of the second moments and their estimated sampling variation. Based on an evaluation of the small sample properties of the OMD estimator, the equally weighted minimum distance (EMD) estimator and an alternative, independently weighted OMD (IWOMD) estimator, which has the same large sample properties as OMD but is unbiased in small samples, Altonji and Segal conclude that, although IWOMD is preferred to OMD, EMD is usually the best of the three estimators they considered. However, EMD estimation of variance components models in levels is not scale invariant. For this reason, I adopt an inverse-variance diagonal weight matrix. Pischke (1995) also found this inverse-variance weighting scheme performed well.

[20] By comparison, the goodness-of-fit statistic of the more restricted model represented by equations (3a) and (3b) with AR(1) transitory error components is 146.91 (p-value<0.001). However, the main reason for the poor fit of this specification appears to be the absence of purely transitory components to account for random measurement error, rather than differential scale factors on the permanent and transitory components: e.g., when purely transitory components are added to the model, the goodness-of-fit statistic is 106.58 (p-value=0.04).

Next, the scale factors on the permanent and transitory components of male wages imply that the variance of each increased roughly equally over the period. For example, the permanent variance increased about 50 percent by 1985, while the transitory variance increased some 30 percent, although there is year-to-year fluctuation in the trends. In contrast, the scale factors on female permanent wages show little increase, while the transitory scale factors increased dramatically relative to 1979, implying that most of the increase in the variance of female wages is due to transitory factors. However, this result should be treated with caution, as females' transitory variance and scale factor parameters are imprecisely estimated. Finally, the intrafamily wage correlations find there is quite strong positive assortive matching of couples (correlation=0.57), and that the correlation between their labor market shocks is relatively weak (0.15), although statistically significant. These estimates confirm that the vast majority (over 90 percent) of the cross-correlation between spouses wages reported in table 2 is due to the permanent correlation.

A Model of Family Labor Supply

Recall that the comparison of the covariance structures of spouses' wages and earnings implies that spouses' hours of work are not independent of wages. I now present a labor supply model which captures the salient feature in the data that the difference between the correlation structures of wages and earnings is due to permanent factors. There are several approach to modelling labor supply behavior, however most studies either model the behavior of males independently of their wives (e.g., Altonji, 1986, and MaCurdy, 1981 and 1985), or model married women's behavior treating their husbands' decisions as exogenous (e.g., Heckman and MaCurdy, 1980, and Mroz, 1987). In contrast, the approach here is to model the labor supply decisions of husbands and wives together, and derive equations for spouses' earnings that depend only on their contemporaneous own-wage, and a time-invariant family effect which summarizes both intra-family and intertemporal behavior.

First, I assume that wages are exogenous with respect to labor supply decisions, and evolve according to equations (5a) and (5b). Second, suppose preferences are homogeneous across families and can be represented by an augmented Stone-Geary utility function, which is additive over time and across spouses' labor supply choices and joint consumption,[21]

$$v(h_{mi},h_{fi},c_i) = \sum_{t=1}^{L} \frac{1}{(1+\delta)^t}\{B_m \ln(\gamma_m - h_{mit}) + B_f \ln(\gamma_f - h_{fit}) + B_c \ln(c_{it} - \gamma_c)\} \tag{6}$$

where families live for a fixed lifetime L; $h_{ji} = (h_{ji1}, \ldots ,h_{jiL})$, $(j=m,\ f)$ are the vectors of lifetime hours of work of the husband and wife in family i; $c = (c_{i1}, \ldots ,c_{iL})$ is the vector of lifetime joint consumption; γ_m and γ_f represent the maximum feasible hours of work for males and females respectively, γ_c is minimum necessary family consumption; B_m, B_f, B_c are the relative weights accorded to the labor-leisure and joint consumption choices in the family's utility function; and δ is the rate of time preference.

Third, if couples have perfect foresight with respect to their vectors of lifetime wages, family non-labor income and prices, the family maximizes the lifetime utility function (6) subject to the intertemporal budget constraint

$$\sum_{t=1}^{L} \frac{1}{(1+r)^t}\left\{w_{mit}h_{mit} + w_{fit}h_{fit} + y_{it}^{N} - p_t c_{it}\right\} = 0 \tag{7}$$

where w_{mit} and w_{fit} are the husband's and wife's period t <u>true</u> wages (i.e. $w_{jit} = w_{jit}^{*} - \eta_{jit}$); p_t is the price of goods; y_{it}^{N} is non-labor family income; and r is the real interest rate. Solving the first order conditions, and assuming for the time being that interior solutions exist for each spouse's hours of work, leads to equations for each spouse's earnings which depend linearly on their own contemporaneous wage and a

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[21] Ashenfelter and Ham (1979) and Johnson and Pencavel (1984) discuss Stone-Geary frameworks in the context of analyzing of male labor supply behavior.

common family effect, F_i:

(8a) $$y_{mit} \equiv w_{mit}h_{mit} = \gamma_m w_{mit} - \psi^t \beta_m F_i + \epsilon_{mit}, \text{ and}$$

(8b) $$y_{fit} \equiv w_{fit}h_{fit} = \gamma_f w_{fit} - \psi^t \beta_f F_i + \epsilon_{fit},$$

where $F_i = \sum_{s=1}^{L} \frac{1}{(1+r)^s}\left\{w_{mis}\gamma_m + w_{fis}\gamma_f + y_{is}^N - p_s\gamma_c\right\}$ is the present discounted value of the family's

future discretionary income; $\psi = \frac{1+r}{1+\delta}$; $\beta_j = \frac{B_j}{B_m+B_f+B_c}\left(\sum_{s=1}^{L}\frac{1}{(1+\delta)^s}\right)^{-1}$, (j=m,f); and ϵ_{mit} and ϵ_{fit}

capture measurement error in earnings and/or taste shocks, etc. In what follows, I assume that $r=\delta$ so that $\psi=1$.

Labor supply behavior is captured by the extent to which β_m and β_f are non-zero. That is, if $\beta_m=\beta_f=0$ then equations (8a) and (8b) are simply the "no labor supply" earnings equations (4a) and (4b) with the addition of measurement errors ϵ_{mit} and ϵ_{fit}, and γ_m and γ_f will be equal to the annual hours of work for males and females respectively. On the other hand, if individuals respond positively to the incentives contained in their own wages, γ_m and γ_f will exceed their hours worked, and β_m and β_f will be non-zero, with relative magnitudes which depend on the relative weights accorded each spouse's leisure choice, B_m and B_f, in the utility function (6). Spouses' intertemporal labor supply behavior is identified from variation in their own wages over time. That is, individuals' in families with higher permanent income (F_i) will work, on average, fewer hours; while, intertemporally, they will work relatively more hours when their own wage is higher.

Given the parsimonious specification of the model, it is worth discussing the restrictions it imposes. First, the assumption of perfect foresight means that all wage changes over the period are fully anticipated, so that any labor supply response will reflect a purely intertemporal substitution response along the λ-*constant* (i.e.constant marginal utility of lifetime income) intertemporal labor supply function. Thus, the model is quite restrictive in its ability to interpret behavioral responses to unanticipated wage changes. Second, the utility function (8) constrains a family's demands for leisure in different periods and across spouses to be substitutes in a very particular way. Specifically, individuals' current period labor supply decisions do not depend on their spouse's wages or their own non-contemporaneous wages except through permanent family income. Third, the model assumes homogeneous preferences across families. Although individual labor supply responses differ by family income resources, which may adequately capture differences across the income distribution, the model does not permit any differences according to, for instance, fertility choices or the stage of the lifecycle observed during the sample period. Fourth, the model restricts individuals (λ-*constant*) compensated labor supply responses to their own permanent and transitory wages to be the same, although it does permit differential uncompensated responses through the permanent family income term, F_i, as permanent and transitory wage changes will affect F_i differently. Given the intertemporal nature of the model and the assumption of perfect foresight, it seems reasonable that the compensated response should be independent of whether it is a transitory or permanent change. Fifth, $\psi=1$ restricts the response to the discretionary family income to be constant across the period. Finally, the model assumes that individuals' hours worked represent labor supply choices and are not constrained by the demand side. If labor market shocks generate positively correlated employment and wage shocks, so that individuals hours are constrained following a negative wage or income shock, then the observed response will tend to overstate the labor supply response.

Notwithstanding these limitations, given the ability of the model to capture a common family effect in spouses' labor supply decisions, together with its tractability, it is a useful model to analyze the link between wage and earnings inequality. The labor supply parameters of interest can be estimated by fitting the model for wages and earnings -- equations (5a & b) and (8a & b) -- to the empirical covariance structure of wages and earnings. The discretionary family income will be a function of each spouse's lifetime vector of wages; I make the simplifying assumption that F_i is correlated with each spouse's permanent wages, over the sample period, but uncorrelated with their transitory wages.[22] The labor supply parameters (β_m, β_f, γ_m, γ_f, σ_F^2=Var(F_i), ρ_{mF}=Corr(F_i,α_{mi}) and ρ_{fF}=Corr(F_i,α_{fi})) are identified from the covariance structure of spouses' wages and earnings subject to two additional restrictions: β_m, β_f and σ_F^2 are only identified up to scale; and the signs of β_m, β_f, ρ_{mF} and ρ_{fF} are not separately identified, although each of these parameters are expected to be positive. I normalize $\sigma_F^2=1$ and restrict the sign of ρ_{mF} to be positive.

Table 4 presents estimates of the labor supply parameters for a variety of specifications. The first column contains the results of fitting the model to the 28 means and 406 covariances of spouses' wages and earnings for the sample of continuous participants. Evaluated at the sample means of annual hours worked (2162 for males and 1608 for females), the scale parameters γ_m and γ_f imply λ-*constant* compensated labor supply elasticities of 0.03 for males and 0.44 for females. The estimated permanent family income effects (β_m and β_f) are each positive and statistically significant. The relative magnitudes of these parameters indicate that the income effect is substantially stronger in female labor supply behavior than in male labor supply, and that female leisure choices receive greater weight in the utility function. Likewise, the correlations between permanent discretionary income and the permanent wage components (ρ_{mF} and ρ_{fF}) are positive and statistically significant; curiously, the correlation is stronger for wives than husbands. Also, the parameter estimates for the error components model of wages (not shown) are similar to those in table 3, except that the scale factors on the permanent and transitory components are more precisely estimated, and the autocorrelation in transitory wages is smaller. The estimate of μ_F (1.466) implies that the mean present value of discretionary permanent family income is

[22] For example, if the interest rate r=0 and f_{jt}=1, for all j and t, corr(u_{jit},F_i)=0 would result from the identifying assumption that the sum of the transitory error terms over the sample is zero.

nearly \$15,000 per year. The consistency of the scale effects can be tested: i.e. $\bar{y}_{mt} = \gamma_m\Gamma_{mt} - \beta_m\mu_F$ and $\bar{y}_{ft} = \gamma_f\Gamma_{ft} - \beta_f\mu_F$. The χ^2 statistic for this test is 20.2 (p-value = 0.09), which provides limited support for the scale restrictions imposed by the model.

The GMM goodness-of-fit statistic (1243, with 377 degrees of freedom) indicates that the model provides a poor statistical fit to the data. However, the fit is at least comparable to other models fit to covariance structures in the literature (e.g., Ashenfelter and Card, 1985). Also, although the wage model is relatively unrestricted, the labor supply model used here is extremely parsimonious, so it is not surprising that it fails a general specification test. A more instructive approach to evaluate the model's adequacy may be to test it against specific alternatives of interest. For this purpose I test, first, the "no labor supply" hypothesis that $\beta_m=\beta_f=0$, and second, the extent to which the separability of spouses' leisure choices implied by the model is consistent with the data. The assumption of no labor supply behavior in spouses' hours of work decisions is strongly rejected (χ^2 statistic=44.07, p-value<0.001).[23]

Second, the model's separability restrictions imply that an individual's non-contemporaneous wages and their spouse's wages only affect their current labor supply decision via permanent discretionary income, F_i. One simple way to test the separability across spouses' is to examine whether the contemporaneous wage of an individual's spouse affects their own labor supply decisions in equations (8a) and (8b). That is, I test $\gamma_{mf}=\gamma_{fm}=0$ in the earnings specification: $y_{jit} = \gamma_j w_{jit} - \beta_j F_i + \gamma_{jk} w_{kit} + \epsilon_{jit}$; j,k = m,f, k≠j. A Lagrange Multiplier test rejects this separability hypothesis (χ^2 statistic=14.9, p-value<0.001). However, when the cross-wage terms are included, their effect on the results is quantitatively small: the female wage effect on male earnings is positive but not statistically significant, while the male wage coefficient on female earnings is -0.03 and statistically significant. The latter result implies that a \$1 increase in male hourly wages reduces female annual earnings by \$300 and labor supply

[23] As ρ_{mF} and ρ_{fF} are unidentified in this framework if $\beta_m = \beta_f = 0$, I also test the joint hypothesis that $\beta_m = \beta_f = \rho_{mF} = \rho_{fF} = 0$; the conclusions are the same in this case.

by 35 annual hours: corresponding to an elasticity of 0.04 evaluated at the sample means.

Labor Force Participation and Selection Bias

The discussion so far has assumed the utility maximization problem leads to an interior solution and positive hours of work for both spouses in each year. Ignoring the participation decision, particularly for women, creates a potentially serious selection bias problem. For example, the average annual participation rate for women in the sample is 70 percent, while 51 percent of women <u>do not</u> participate in at least 1 year during the sample period (the corresponding figures for men are 93 percent and 14 percent). In this section I consider two alternative corrections for the selection bias problem associated with female non-participation, which distinguish only between continual and non-continual participants. The first is a structural selection model derived from the first order conditions for the labor supply equation for females (equation (8b)). This method is similar to approaches in Ashenfelter and Card (1985), and Heckman and Robb (1985) in the context of controlling for non-random selection into training programs, and draws on Lee's (1982, 1983) maximum order statistic approach to selection correction. The second is a simple data driven approach which reweights the sample using observed covariates to predict the probability of selection. This approach is analogous to that of reweighting samples to take account of non-random attrition using observable characteristics (e.g., Fitzgerald, Gottschalk, and Moffitt, 1998).

First, the female earnings equation (8b) implies that a woman will participate in year t if $h_{fit} > 0$, or if her wage $w_{fit} > (\beta_f/\gamma_f)F_i$. Expressed in terms of the error components specification for wages, she will participate in <u>each</u> of the T sample years if

$$(9) \qquad \min_{1 \leq s \leq T}\{w_{fis}\} = \min_{1 \leq s \leq T}\{\Gamma_{fs} + \phi_{fs}u_{fis}\} + \min_{1 \leq s \leq T}\{\theta_{fs}\}\alpha_{fi} + \xi_i^* > (\beta_f/\gamma_f)F_i$$

where ξ_i^* is a non-negative random variable. In this selection decision, wages have two roles: first, the substitution effect implies that higher wages will increase the probability of participating; however, higher (permanent) wages also raise the family's discretionary income, hence the reservation wage, and thus reduces the probability of participating. This suggests that non-participants during the sample period will include both those with relatively low permanent wages (compared to F_i), and those with relatively low transitory wages.[24]

In order to obtain a tractable empirical selection specification, I simplify the selection rule (9) and assume that participation occurs in each year if and only if

$$(10) \qquad z_i = (\beta_f/\gamma_f)F_i - \alpha_{fi} + \xi_i < z_0$$

where z_0 is constant across families, and ξ_i is a random variable, uncorrelated with F_i and α_{fi}, which captures individual-specific transitory wage and taste shocks over the sample period.[25] To proceed, I also assume that all the random variables in the model are joint normally distributed.

The moments, conditional on sample selection, are directly comparable with the observed sample moments. The procedure here is to use the selection rule (10), together with the wage and earnings equations (5a & b) and (8a & b) to describe the means and covariances of earnings and wages of the population from which the sample is selected. Under the normality assumption, the selection bias effect is characterized by two parameters. The sample means differ from the population means by the product of a parameter λ_1 and the covariance between the variable of interest and the selection variable, where $\lambda_1 = E(z_i|z_i<z_0)/Var(z_i)$. Similarly, the sample covariances differ from the population covariances by

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[24] As the sample selection problem arises from <u>any</u> non-participation, we make no distinction according to the number of years of non-participation. However, one possible way to distinguish these two types of participants would be according to the frequency of non-participation. For example, frequent non-participants are likely to be those with relatively low permanent wages; while infrequent non-participants are likely to be those with relatively low transitory wages.

[25] The results in table 3 suggest that θ_{ft} lies in the range 1 to 1.2, and $\theta_{ft}\alpha_{fi}$ is minimized in year 1 if $\alpha_{fi}>0$, and in year 7 if $\alpha_{fi}<0$; as a simplification I ignore the time variation in the scale factors. Also, z_0 can be interpreted as representing the "average" wage level, while ξ_i captures the combined effect of transitory wages and ξ_i^* in equation (9).

the product of a parameter λ_2 and the covariances between each of the variables of interest and the selection variable, where $\lambda_2 = (v\text{-}1)/Var(z_i)^2$, and v is the variance of a standard normal variate truncated at $z_0/Var(z_i)^{1/2}$.[26] This specification does not allow separate identification of the 14 wage means, the mean of F_i and the selection bias parameter λ_1: for identification, I restrict the first year population mean of male wages to equal the sample mean.[27]

The results from this exercise are presented in column (2) of table 4. The estimates of λ_1 and λ_2 are small and insignificantly different from 0, implying there is no evidence of selection bias on either the first or second moments of the data. On the face of it, the absence of selection bias is somewhat surprising. For example, although average wages and hours of males is quite similar across various sample selection criteria (see appendix table A1), female wages and hours are about 10 percent higher, and the variance of earnings is about 20 percent higher, in the sample of women who are continuous participants than in the sample which also includes non-participants. However, if non-participation is determined mainly by low transitory, rather than permanent, wages the sample selection may be roughly random and result in little bias. Given the absence of selection bias, it is unsurprising that the parameter estimates for the earnings equations are close to those presented in column (1), although the LM statistic easily rejects the consistency of the first moment scale factors in this specification, and the cross-wage effects are also stronger. One possible explanation for these results may be that the restriction used to identify the first moment parameters is not valid. However, the conclusion of insignificant second moment selection bias in this specification is robust to the validity of this restriction.

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[26] Specifically, the conditional means of wages and earnings are $E(w_{jit}|z_i<z_0) = E(w_{jit}) + Cov(w_{jit},z_i)\lambda_1$, and $E(y_{jit}|z_i<z_0) = E(y_{jit}) + Cov(y_{jit},z_i)\lambda_1$; and the conditional covariances are $Cov(w_{jit},w_{kis}|z_i<z_0) = Cov(w_{jit},w_{kis}) + Cov(w_{jit},z_i)Cov(w_{kis},z_i)\lambda_2$, $Cov(y_{jit},y_{kis}|z_i<z_0) = Cov(y_{jit},y_{kis}) + Cov(y_{jit},z_i)Cov(y_{kis},z_i)\lambda_2$, and $Cov(w_{jit},y_{kis}|z_i<z_0) = Cov(w_{jit},y_{kis}) + Cov(w_{jit},z_i)Cov(y_{kis},z_i)\lambda_2$.

[27] This restriction will be valid if male wages are unaffected by the participation decisions of their wives, which appears to be approximately true in table A1. The remaining parameters of interest are identified by the second moments of the data and unaffected by this restriction.

One problem with this structural selection model is that it relies on functional form assumptions for identification and may have little power to detect non random selection effects. With this caveat in mind, I consider an alternative selection correction which reweights the data to account for non-randomness in female participation decisions. The comparative strength of this approach is that it does not rely on functional form and distributional assumptions for identification; while its comparative weakness is that it does not account for selection on unobservables. I first fit a Probit selection model for wives' continuous-participation to the sample of 1830 couples with continuously employed males (described in column 4 of table 1). The set of covariates used to predict selection are quadratics in the wife's 1980 age, the number of children and the age of the youngest child; the wife's years of education; and dummy variables for whether the wife is Black, in the SEO subsample, the presence of children under 6, the 1980 region of residence and whether the region of residence changed during the period. The results are presented in appendix table A4, and suggest that Black and educated women are more likely to participate in all years, while women with children, in their middle age, or who changed region of residence are less likely to be continuous participants. The model has a pseudo-R^2=0.07, while the average predicted selection probability is 0.53 in the selected subsample, compared to 0.44 in the non-selected subsample.

In the second stage, each observation in the narrower sample of 890 continuously employed couples is reweighted by its predicted selection probability, and the joint wage and labor supply model is re-estimated using the reweighted first and second moments. The results from this approach are summarized in column (3) of table 4. Again, the estimates using this approach are similar to those in columns (1) and (2). Interestingly, the goodness-of-fit statistic for this specification is substantially lower than for the other specifications, although it still represents a poor statistical fit of the data. Although both of the selection-bias corrections considered here have weaknesses, the results using each are similar to the uncorrected model in column (1). This provides some confidence concerning the relevance of the results to a broader population of families. At the least, the results can be interpreted as pertaining to the population of continuously married and participating couples.

III: Implications for Earnings Inequality

In this section I use the estimates from the model for spouses' wages and their labor supply decisions to evaluate the implications of the permanent-transitory nature of wage differences, intrafamily wage interactions, and the labor supply decisions of families for understanding individual and family earnings inequality. The framework developed enables a structural interpretation of how much of the levels of, and changes in, cross-sectional and permanent earnings inequality may be explained by intertemporal labor supply behavior and, for family earnings inequality, assortive-mating and correlated labor market shocks of spouses. Equations (5a & b) and (8a & b) imply that the cross-sectional variance of spouses' individual and joint family earnings can be decomposed into components attributable to the permanent and transitory wages, family permanent income F_i, and random effects.[28] The terms involving permanent wages and family income determine the level of permanent inequality.

Table 5 presents the implications for earnings inequality using the models estimated in the previous section. In order to get a sense of how well the framework does in terms of matching the level and change in earnings inequality in the sample, the first set of rows compare the average actual level of inequality of the sample period and the change in inequality between 1979 and 1985, with the predictions using the model in table 4, column (1) which does not correct for sample selection. The results for the levels are very close, while for the specific year-to-year changes they predictions are less accurate, particularly for female inequality. The second set of rows in table 5 presents the predicted average cross-

[28] That is, $\mathrm{Var}(y_{jit}) = \gamma_j^2\theta_{jt}^2 \mathrm{var}(\alpha_{ji}) + \beta_j^2\sigma_F^2 - 2\beta_j\gamma_j\theta_{jt}\mathrm{cov}(\alpha_{ji},F_i) + \gamma_j^2\phi_{jt}^2\mathrm{var}(u_{jit}) + \mathrm{var}(\epsilon_{jit})$, $(j=m,f)$; and $\mathrm{Var}(y_{mit}+y_{fit}) = \gamma_m^2\theta_{mt}^2\mathrm{var}(\alpha_{mi}) + \gamma_f^2\theta_{ft}^2\mathrm{var}(\alpha_{fi}) + 2\gamma_m\gamma_f\theta_{mt}\theta_{ft}\mathrm{cov}(\alpha_{mi},\alpha_{fi}) + (\beta_m+\beta_f)^2\sigma_F^2 - 2(\beta_m+\beta_f)(\gamma_m\theta_{mt}\mathrm{cov}(\alpha_{mi},F_i) + \gamma_f\theta_{ft}\mathrm{cov}(\alpha_{fi},F_i) + \gamma_m^2\phi_{mt}^2\mathrm{var}(u_{mit}) + \gamma_f^2\phi_{ft}^2\mathrm{var}(u_{fit}) + 2\gamma_m\gamma_f\phi_{mt}\phi_{ft}\mathrm{cov}(u_{mit},u_{fit}) + \mathrm{var}(\epsilon_{mit}) + \mathrm{var}(\epsilon_{fit})$.

sectional inequality and changes over the period, together with the fractions attributable to permanent factors, for the population of continuously married couples with husband working, using the model in table 4, column (2) which corrects for the sample selection of continuously employed wives. This model predicts higher levels of inequality than the uncorrected model, particularly for female earnings, but almost the same changes. Although these results are not directly comparable to those in column (4) of table 1 (because of the unbalanced sample analysis in those results), the implications that the sample selection of employed wives tends to reduce the level of earnings inequality somewhat more than the change appear to be roughly similar. About two-thirds of the cross-sectional inequality in male earnings is due to permanent factors, while over three-quarters of female and family earnings inequality is permanent. The increases in male and family earnings inequality between 1979 and 1985 are due proportionately to permanent and transitory factors, while the more modest increase in female inequality is due mainly due to an increase in transitory inequality.

I now consider the contributions of intrafamily correlations and labor supply behavior to cross-sectional and permanent inequality. This requires the construction of counterfactuals for the nature of earnings inequality in the absence of labor supply behavior, and also in the absence of correlated permanent and transitory wages components. In the absence of labor supply behavior, I assume that husbands' and wives' work fixed numbers of hours equal to the sample average earnings divided by wages implied by the model, and use the wage model parameters to construct the counterfactual predictions. The counterfactual predictions with no labor supply and no intrafamily interactions are constructed by also assuming that the correlations between spouses' permanent and transitory wages are zero.

The third set of rows in table 5 summarize the contributions of within-family wage correlations and labor supply behavior to the levels and changes in cross-sectional earnings inequality over the period. These results show that about 20 percent of the observed level of family earnings inequality, and slightly less of the increase in family earnings inequality between 1979 and 1985, is due to the combined effects of positive assortive matching and correlated labor market shocks on husbands and wives' wages. That is, holding the level of individual earnings inequality constant, but randomly matching spouses' would lower family inequality by 20 percent. As expected given the modest estimates of male labor supply response in table 4, labor supply behavior contributes only 4 percent of the level and 8 percent of the change in male earnings inequality. In contrast, labor supply behavior accounts for over one-third of female earnings inequality over the period, and nearly two-thirds of the increase. The corresponding estimates of the labor supply contributions to family earnings inequality are 11 and 10 percent.

Finally, the fourth set of rows in table 5 presents corresponding estimates of the assortive-mating and labor-supply contributions to permanent inequality. Assortive matching contributes over one-quarter of the level of permanent inequality, and 23 percent of the increase. The estimates of the labor supply contributions to permanent inequality are generally similar to their respective contributions to cross-sectional inequality. The absence of any significant male labor supply behavior implies it has little effect on the transmission of wage inequality into earnings inequality, and that changes in earnings inequality will simply reflect changes in wage inequality. However, labor supply behavior appears to play an important role in understanding female earnings inequality, and a more modest role in family earnings inequality over the 1979-85 period.

IV: Conclusion

This paper has provided an analysis of the transmission of wage inequality into earnings inequality, focusing on the extent of permanence in wage differences, the nature of intrafamily correlations, and the role of labor supply behavior. I conclude that the increase in male and family earnings inequality during the early 1980s is due roughly proportionately to permanent and transitory factors, while the modest increase in female earnings inequality appears to be due mainly to an increase in transitory factors. Spouses' positively correlated wage shocks explain about 20 percent of the overall increase in family earnings inequality, and positive assortive mating about 25 percent of the increase in permanent inequality. Also, although male labor supply behavior plays a relatively minor role in understanding earnings inequality, the behavior of wives contributes substantially to the increases in female and family earnings inequality over the period.

However, there are some caveats to the analysis. Most importantly perhaps is that the model assumed that the changing opportunities that individuals faced over the period were perfectly foreseen, so that their responses represented purely intertemporal substitutions along a constant marginal utility labor supply function. This does not allow for any behavioral responses to unanticipated income effects, which is an obvious drawback given the extent of non-stationarity over the period. Assuming such unanticipated changes are viewed largely as income shocks then, arguably, the behavioral responses would tend to mitigate against the increase in inequality.

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Appendix A: Sample Selection and Characteristics

The sample selection criteria are that couples be continuously married between 1979 and 1985; both husband and wife be aged between 18 and 60 in 1980; that annual earnings not be top-coded in any year; and that individuals' wages were less than $100 per hour in all years. In nominal terms, male annual earnings in 1979-81 are top-coded at $99,999, and in 1982-85 at $999,999; female earnings are top-coded at $99,999 throughout the entire sample period. Top-coded observations have two adverse effects on the analysis: the reported earnings of those individuals affected will be largely static due to serial correlation in earnings; and, as a result of the change in the male top-coding value in 1982, when top-coded observations are included, the variance of reported male earnings more than doubles from 1981 to 1982. Average hourly wages are measured as annual earnings divided by annual hours worked. Both earnings and wages are expressed in constant (1987) dollars, obtained by deflating nominal values by the Consumer Price Index.

The first sample includes all couples who satisfy these criteria. The number of observations affected by top-coding was 86, and 15 observations had wages > $100 in any year. An additional 28 observations were excluded as *outliers* because of their effect on the variance of wages or earnings in particular years, while the autocovariances are largely unaffected when these observations are excluded. (Because the analysis is in levels, a very small number of observations can have a large effect on the second moments of the data.) The second sample excludes couples if the male does not report positive earnings and hours in every year of the sample. The third sample, which is primarily used in this paper, also excludes couples if the female does not report positive earnings and hours in each year of the sample.

Table 1: Earnings Inequality: Married Couples, 1979-85

	CPS Data				PSID Data					
	Aged 18-60 in Survey Year		Aged 18-60 in 1980		Continuously Married		Husband Participating		Both Spouses Participating	
	(1)		(2)		(3)		(4)		(5)	
Male earnings										
1979	28.7	0.30	28.7	0.30	25.9	0.23	26.5	0.21	25.9	0.20
1982	26.4	0.40	26.7	0.40	24.7	0.31	25.5	0.28	25.0	0.26
1985	28.1	0.42	28.9	0.41	26.9	0.34	27.7	0.31	27.2	0.28
% Change (1979-85)	-2.3	38.2	0.7	35.8	3.9	46.4	4.6	45.1	4.9	35.8
Female earnings										
1979	11.1	0.63	11.1	0.63	11.0	0.56	11.1	0.56	12.8	0.40
1982	11.4	0.62	11.5	0.61	10.7	0.62	10.8	0.63	13.0	0.42
1985	12.6	0.68	13.0	0.65	11.9	0.62	12.0	0.61	14.7	0.41
% Change (1979-85)	12.8	7.3	16.5	2.9	8.5	10.4	8.4	9.2	15.2	0.7
Family earnings										
1979	33.9	0.31	33.9	0.31	33.2	0.23	34.5	0.20	38.7	0.16
1982	31.5	0.39	31.7	0.39	31.3	0.29	33.0	0.25	38.0	0.20
1985	34.5	0.41	35.0	0.41	34.3	0.32	36.5	0.27	41.9	0.20
% Change (1979-85)	1.7	32.1	3.2	32.7	3.3	44.1	5.7	36.4	8.3	23.9
Sample Sizes										
1979	31692		31692		2143		1830		890	
1982	27819		28878							
1985	26479		27607							

Notes: The first entry for each sample is mean earnings (in 1987 $000); the second entry is the squared coefficient of variation of earnings. Columns (1) and (2) use data from the 1980, 1983 and 1986 March Current Population Surveys (CPS), and pertain to the previous calender year. Columns (3) -- (5) use data from the 1986 release of the Panel Study of Income Dynamics (PSID). All calculations are conditional on positive earnings reported.

Table 2: Summary of the Covariance Structures of Spouses' Wages and Earnings

	Average Variance	Average Correlation at lag k 0	1	2	3	4	5	6
	A: Correlation Structure of Intrafamily Wages							
Between Wages of								
Males in t, t+k	36.918		0.769	0.727	0.704	0.650	0.615	0.624
Females in t, t+k	22.113		0.734	0.683	0.667	0.645	0.636	0.629
Males in t, Females in t+k		0.340	0.319	0.319	0.327	0.313	0.304	0.328
Males in t+k, Females in t		0.340	0.324	0.322	0.327	0.312	0.317	0.363
	B: Correlation Structure of Intrafamily Earnings							
Between Earnings of								
Males in t, t+k	1.647		0.834	0.758	0.719	0.671	0.628	0.590
Females in t, t+k	0.727		0.873	0.817	0.776	0.745	0.717	0.688
Families in t, t+k	2.909		0.870	0.818	0.794	0.761	0.734	0.709
Males in t, Females in t+k		0.246	0.242	0.249	0.259	0.264	0.270	0.261
Males in t+k, Females in t		0.246	0.244	0.248	0.258	0.255	0.254	0.271

Notes: The full covariance matrices for wages and earnings are presented in appendix tables A2 and A3 respectively. Earnings have been divided by 10,000.

Table 3: Estimates of Error Components Model for Wages

Permanent Components		Transitory Components	
$\mathrm{Var}(\alpha_{mi})$	16.204	$\mathrm{Var}(\upsilon_{mit})$	3.244
	(3.06)		(0.93)
		ρ_m	0.822
			(0.07)
$\mathrm{Var}(\alpha_{fi})$	13.171	$\mathrm{Var}(\upsilon_{fit})$	0.489
	(2.40)		(0.47)
		ρ_f	0.843
			(0.08)
$\mathrm{Corr}(\alpha_{mi},\alpha_{fi})$	0.572	$\mathrm{Corr}(\upsilon_{mit},\upsilon_{fit})$	0.154
	(0.05)		(0.06)
θ_{m80}	0.998	ϕ_{m80}	0.962
	(0.08)		(0.10)
θ_{m81}	0.961	ϕ_{m81}	1.151
	(0.07)		(0.13)
θ_{m82}	1.086	ϕ_{m82}	1.030
	(0.08)		(0.16)
θ_{m83}	1.164	ϕ_{m83}	0.956
	(0.10)		(0.19)
θ_{m84}	1.034	ϕ_{m84}	1.351
	(0.11)		(0.27)
θ_{m85}	1.244	ϕ_{m85}	1.148
	(0.11)		(0.18)
θ_{f80}	0.906	ϕ_{f80}	1.563
	(0.08)		(0.65)
θ_{f81}	0.857	ϕ_{f81}	2.192
	(0.11)		(1.11)
θ_{f82}	0.905	ϕ_{f82}	1.773
	(0.09)		(0.90)
θ_{f83}	0.893	ϕ_{f83}	1.900
	(0.09)		(0.97)
θ_{f84}	0.890	ϕ_{f84}	2.424
	(0.11)		(1.33)
θ_{f85}	1.028	ϕ_{f85}	1.967
	(0.12)		(1.12)
		$\mathrm{Var}(\eta_{mit})$	4.973
			(0.69)
		$\mathrm{Var}(\eta_{fit})$	4.614
			(0.49)
Goodness-of-fit	88.95 (71)		

Notes: Estimated standard errors in parentheses, except χ^2 degrees of freedom for Goodness-of-fit statistic. Estimation by Minimum Distance Estimation using Inverse Variance Weights. Number of observations are 890 couples over 7 years.

Table 4: Estimates of Labour Supply Model Parameters

	(1)	(2)	(3)
Parameters:			
γ_m	0.223 (0.01)	0.222 (0.01)	0.222 (0.01)
γ_f	0.233 (0.02)	0.230 (0.02)	0.225 (0.02)
μ_F	1.466 (0.30)	1.611 (0.61)	1.430 (0.36)
β_m	0.099 (0.03)	0.076 (0.07)	0.087 (0.03)
β_f	0.420 (0.07)	0.368 (0.10)	0.372 (0.06)
ρ_{mF}	0.429 (0.07)	0.439 (0.10)	0.403 (0.08)
ρ_{fF}	0.552 (0.16)	0.482 (0.18)	0.469 (0.19)
λ_1	.---	-0.008 (0.04)	.---
λ_2	.---	-0.005 (0.01)	.---
Specification Tests:			
Goodness-of-fit	1242.70 (377)	1242.87 (376)	1119.18 (377)
No Labor supply ($\beta_m=\beta_f=0$)	44.07 (2)	22.24 (2)	59.11 (2)
Cross-wage effects	14.87 (2)	44.89 (2)	19.21 (2)
Consistency of scale effects	20.17 (13)	35.51 (12)	19.59 (13)
Selection bias	.---	0.39 (2)	.---

Notes: Estimation is by Minimum Distance using Inverse Variance Weights. Number of observations are 890 couples observed over 7 years. Estimated standard errors in parentheses after parameters; χ^2 degrees of freedom in parentheses after test statistics. Earnings have been divided by 10,000. The No Labor supply and selection bias tests are based on Wald statistics; cross-wage and consistency of scale effects tests are based on LM statistics. In column (1), the model is estimated ignoring sample selection issues; in column (2), the model corrects for selection bias using a structural selection approach for women's participation decisions; in column (3), the model corrects for selection bias by reweighting the data to account for women's participation decisions.

Table 5: Implications for Earnings Inequality

	Earnings of		
	Males	Females	Families
1. Sample predictions:			
Average actual CV^2	0.244	0.400	0.187
Average predicted CV^2	0.241	0.400	0.186
1979-85 actual change	0.073	0.003	0.039
1979-85 predicted change	0.062	0.034	0.038
2. Population predictions:			
Average CV^2	0.245	0.452	0.199
Fraction permanent	0.677	0.779	0.773
1979-85 change	0.063	0.036	0.041
Fraction permanent	0.713	0.326	0.762
3. Fraction of CV^2 due to:			
Intrafamily correlations	---	---	0.211
Labor supply	0.042	0.374	0.111
Fraction of 1979-85 change due to:			
Intrafamily correlations	---	---	0.168
Labor supply	0.078	0.654	0.212
4. Fraction of permanent CV^2 due to:			
Assortive mating	---	---	0.277
Labor supply	0.033	0.383	0.106
Fraction of 1979-85 change due to:			
Assortive mating	---	---	0.230
Labor supply	0.068	0.568	0.185

Notes: The "sample predictions" are based on the model presented in table 4, column (1), which ignores sample selection bias. The "population predictions" are based on the model presented table 4 column (2), which adjusts for sample selection bias.

Table A1: Sample Characteristics

	CPS Families	PSID Families		
	Married Couples 1979-85	Married Couples 1979-85	Husband Working 1979-85	Both Spouses Working 1979-85
Sample Size, N	---	2143	1830	890
Annual hours worked:[a] Males	2117	2147	2180	2162
Females	1503	1435	1429	1608
Hourly wage rate:[a] Males	13.28	12.20	12.43	12.29
Females	7.78	7.57	7.66	8.41
Age (1980): Males	39.86	37.06	35.70	36.06
Females	37.29	34.49	33.34	33.69
Number of Children (1980)	1.27	1.48	1.49	1.41
Proportion Black: Males	0.07	0.26	0.24	0.26
Females	0.07	0.24	0.23	0.26
Education:[b] Males	12.74	12.73	13.01	13.21
Females	12.46	12.67	12.83	13.21
Average participation, 1979-85				
Males	0.82	0.93	-.-	-.-
Females	0.62	0.70	0.72	-.-
Proportion SEO Subsample	-.-	0.37	0.35	0.35

Notes: Sample selection criteria also require males and females aged 18-60 in 1980, non top-coded earnings, wages < $100/hour.

(a) Means of annual hours worked and hourly wages are calculated conditional on positive hours and earnings. Wages are in $1987.

(b) Education in column (1) is recorded as years of completed education in the 1986 CPS. Education in the PSID is measured as the highest level reported in the 1980-86 surveys using the following categorical scheme: 1 = '0-5 grades'; 2 = '6-8 grades'; 3 = '9-11 grades'; 4 = '12 grades'; 5 = '12 plus nonacademic training'; 6 = 'some college'; 7 = 'college degree, not advanced'; 8 = 'college and advanced degree'. Years of education are imputed at the midpoints of categories 1 -- 4, and 13, 14, 16, and 18 respectively for categories 5 -- 8.

Table A2: Empirical Covariance Matrix: Husbands' and Wives' Wages, 1979-85

And	Males in 1979	Males in 1980	Males in 1981	Males in 1982	Males in 1983	Males in 1984	Males in 1985	Females in 1979	Females in 1980	Females in 1981	Females in 1982	Females in 1983	Females in 1984	Females in 1985
Males in														
1979	30.909	0.805	0.728	0.742	0.679	0.572	0.624	0.357	0.342	0.287	0.347	0.322	0.278	0.328
	(1.70)													
1980	24.413	29.766	0.788	0.787	0.742	0.596	0.658	0.364	0.351	0.304	0.330	0.322	0.296	0.330
	(1.52)	(1.69)												
1981	23.266	24.724	33.049	0.802	0.735	0.610	0.675	0.329	0.304	0.295	0.325	0.312	0.291	0.321
	(1.53)	(1.57)	(2.04)											
1982	24.131	25.129	26.972	34.238	0.838	0.644	0.723	0.366	0.340	0.304	0.352	0.326	0.302	0.347
	(1.64)	(1.63)	(1.74)	(2.11)										
1983	22.894	24.560	25.610	29.715	36.760	0.678	0.741	0.353	0.346	0.322	0.359	0.348	0.320	0.365
	(1.70)	(1.65)	(1.70)	(1.85)	(2.34)									
1984	22.346	22.845	24.647	26.474	28.873	49.385	0.704	0.304	0.264	0.261	0.284	0.279	0.292	0.299
	(1.69)	(1.53)	(1.68)	(1.68)	(1.85)	(6.56)								
1985	23.100	23.895	25.833	28.170	29.927	32.952	44.318	0.363	0.330	0.319	0.335	0.336	0.336	0.387
	(1.76)	(1.71)	(1.77)	(1.88)	(2.03)	(2.32)	(2.82)							
Females in														
1979	8.906	8.907	8.474	9.590	9.587	9.582	10.829	20.091	0.731	0.618	0.665	0.651	0.597	0.629
	(1.06)	(1.00)	(1.10)	(1.25)	(1.16)	(1.27)	(1.34)	(1.62)						
1980	8.423	8.482	7.735	8.808	9.298	8.222	9.731	14.516	19.611	0.668	0.691	0.685	0.651	0.675
	(0.89)	(0.92)	(0.94)	(1.01)	(1.09)	(1.11)	(1.18)	(1.12)	(1.39)					
1981	7.842	8.164	8.333	8.756	9.587	9.030	10.459	13.629	14.540	24.180	0.737	0.663	0.595	0.634
	(1.09)	(1.24)	(1.32)	(1.24)	(1.21)	(1.38)	(1.59)	(1.29)	(1.22)	(3.32)				
1982	8.695	8.119	8.415	9.268	9.794	8.975	10.054	13.426	13.788	16.316	20.283	0.775	0.692	0.724
	(1.02)	(1.03)	(1.08)	(1.15)	(1.21)	(1.29)	(1.28)	(1.15)	(1.06)	(1.37)	(1.46)			
1983	8.184	8.034	8.197	8.731	9.661	8.968	10.231	13.348	13.885	14.917	15.982	20.947	0.733	0.750
	(1.10)	(1.08)	(1.15)	(1.28)	(1.32)	(1.39)	(1.53)	(1.28)	(1.12)	(1.33)	(1.36)	(1.71)		
1984	7.765	8.102	8.399	8.857	9.738	10.296	11.228	13.441	14.476	14.697	15.640	16.841	25.192	0.762
	(1.12)	(1.07)	(1.21)	(1.14)	(1.32)	(1.69)	(1.66)	(1.24)	(1.11)	(1.42)	(1.27)	(1.34)	(2.04)	
1985	9.025	8.914	9.134	10.047	10.962	10.389	12.734	13.947	14.796	15.426	16.133	16.986	18.937	24.489
	(1.19)	(1.14)	(1.20)	(1.31)	(1.45)	(1.48)	(1.59)	(1.27)	(1.18)	(1.39)	(1.40)	(1.52)	(1.48)	(1.79)

Notes: Covariances are below the diagonal, correlations are above the diagonal, and estimated standard errors of covariances in parentheses. Number of Observations = 890. Wages are in constant $1987. Year-specific means removed.

Table A3: Empirical Covariance Matrix: Husbands' and Wives' Earnings, 1979-85

	Covariance / Correlation between Earnings of													
	Males in							Females in						
And	1979	1980	1981	1982	1983	1984	1985	1979	1980	1981	1982	1983	1984	1985
Males in														
1979	1.362	0.834	0.746	0.711	0.658	0.610	0.590	0.235	0.247	0.249	0.269	0.264	0.269	0.261
	(0.07)													
1980	1.111	1.301	0.832	0.771	0.725	0.663	0.646	0.251	0.274	0.260	0.273	0.270	0.268	0.271
	(0.06)	(0.07)												
1981	1.052	1.147	1.459	0.848	0.769	0.705	0.693	0.245	0.261	0.258	0.263	0.244	0.246	0.261
	(0.06)	(0.07)	(0.09)											
1982	1.055	1.119	1.304	1.618	0.856	0.748	0.736	0.266	0.271	0.248	0.257	0.237	0.240	0.252
	(0.07)	(0.07)	(0.08)	(0.10)										
1983	1.028	1.107	1.244	1.457	1.792	0.798	0.757	0.272	0.283	0.262	0.263	0.241	0.236	0.241
	(0.07)	(0.07)	(0.08)	(0.09)	(0.11)									
1984	0.997	1.059	1.192	1.332	1.496	1.961	0.833	0.262	0.238	0.234	0.231	0.205	0.211	0.209
	(0.07)	0.07)	(0.08)	(0.08)	(0.10)	(0.15)								
1985	0.982	1.051	1.194	1.335	1.445	1.663	2.035	0.271	0.246	0.255	0.248	0.230	0.233	0.249
	(0.07)	(0.07)	(0.08)	(0.08)	(0.10)	(0.12)	(0.14)							
Females in														
1979	0.222	0.232	0.240	0.274	0.295	0.298	0.314	0.657	0.835	0.793	0.745	0.718	0.715	0.688
	(0.04)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)						
1980	0.229	0.248	0.250	0.274	0.301	0.265	0.278	0.538	0.632	0.870	0.803	0.771	0.756	0.719
	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)					
1981	0.239	0.243	0.256	0.259	0.288	0.269	0.298	0.528	0.568	0.674	0.865	0.809	0.797	0.762
	(0.04)	(0.03)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.05)				
1982	0.266	0.264	0.269	0.277	0.299	0.275	0.300	0.512	0.541	0.602	0.719	0.874	0.826	0.790
	(0.04)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.05)			
1983	0.265	0.265	0.253	0.259	0.277	0.246	0.282	0.500	0.526	0.571	0.637	0.738	0.891	0.856
	(0.04)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)		
1984	0.279	0.271	0.264	0.271	0.280	0.262	0.295	0.515	0.533	0.581	0.622	0.680	0.788	0.900
	(0.04)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.05)	(0.04)	(0.05)	
1985	0.285	0.290	0.295	0.301	0.303	0.274	0.333	0.523	0.536	0.586	0.628	0.689	0.749	0.879
	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.05)	(0.06)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.05)

Notes: Covariances are below the diagonal, correlations are above the diagonal, and estimated standard errors of covariances in parentheses. Number of Observations = 890. Earnings are in constant $1987 (divided by 10,000). Year-specific means removed.

Table A4: Sample Selection Equation Estimates

	Dependent Variable: Participation in Each Year = 1
Intercept	-4.624
	(0.85)
Age	0.212
	(0.04)
Age Squared	-0.302
(/100)	(0.06)
Race (Black=1)	0.617
	(0.15)
Years of Schooling	0.157
	(0.02)
Northeast	0.100
	(0.17)
Northcentral	-0.269
	(0.16)
South	0.007
	(0.15)
Change in Region	-0.935
	(0.24)
No. Children	-0.719
	(0.15)
No. Children Squared	0.071
	(0.03)
Children Under 6	-0.209
	(0.38)
Youngest Child's Age	0.115
	(0.07)
Youngest Child's Age Squared	-0.882
(/100)	(0.82)
SEO Sub-sample	-0.086
	(0.13)
Psuedo R^2	0.070
No. Observations	1830
Predictions	
Mean Selection Probability	0.534
(Selected Subsample, N=890)	(0.005)
Mean Selection Probability	0.441
(Non-selected saubsample, N=940)	(0.005)

Notes: Standard errors are in parentheses. The model is estimated as a Probit specification. Age and Race pertain to the woman. The excluded region is West. Change in Region is a dummy variable equal to 1 if the region of residence changed during the sample period.