# HOUSEWORK AND THE WAGES OF YOUNG, MIDDLE-AGED, AND OLDER WORKERS

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*This article uses samples of young, middle-aged, and older married workers drawn from the Panel Study of Income Dynamics to examine whether the effect of housework time on wages differs among these age groups. Results from OLS, fixed effects, and panel data instrumental variables models show that young and middle-aged wives are the only groups for which the authors find consistent evidence of a housework effect on wages. Each additional hour of housework reduces their wages by 0.1–0.4%. Additionally, the analysis finds evidence that for young workers, housework time is an important determinant of the male/female wage gap.* (*JEL* J16, J22, J31)

## I. INTRODUCTION

> If childcare and other housework demand relatively large quantities of "energy" as compared to leisure and other non-market uses of time by men, women with responsibilities for housework would have less energy available for the market than men would. This would reduce the hourly earnings of married women, affect their jobs and occupations, and even lower their investment in market human capital when they work the same number of market hours as married men. Consequently, the housework responsibilities of married women may be the source of much of the difference in earnings and in job segregation between men and women.
>
> —Gary S. Becker (1985, p. S35)

Consistent with Becker's hypothesis, employed women spend more hours on housework per week than do employed men on average. The magnitude of this gender gap in housework time is large, despite a reallocation of some housework activities from wives to husbands in recent years. Over the past 30 years, women's time spent in housework declined from four to three times that of men (Blau et al., 2002, p. 56). Because child care and housework continue to be primarily a woman's responsibility, many researchers have examined the extent to which the gender gap in housework time contributes to the gender wage gap. In general, these studies report that time spent on housework results in wage penalties, especially for married women, and including it in a wage regression increases the explained portion of the gender wage gap.

What has not been examined is the effect of housework time on wages over the life cycle. Many aspects of housework are likely to vary over the life cycle, if for no other reason than individuals' fertility decisions are a function of age. It seems reasonable to assume that the total amount of housework required to maintain a family would depend on the size of the family and the ages of its members. As a young couple starts their family, the amount of housework necessary to take care of its members will increase. Conversely, as children age and leave the home, the total amount of required housework should decline.

**ABBREVIATIONS**

GLS: Generalized Least Squares
HTIV: Hausman-Taylor Instrumental Variables
IV: Instrumental Variables
OLS: Ordinary Least Squares
PSID: Panel Study of Income Dynamics

*This is a revision of a paper presented at the 77th Annual Conference Western Economic Association International June 29–July 3, 2002. This work was started while Keith was a Visiting Scholar at the Survey Research Center of the Institute for Survey Research, University of Michigan, Ann Arbor, Michigan. The authors are grateful to Joni Hersch, Daniel H. Hill, Bradley Lagger, John E. Murray, Leslie S. Stratton, and anonymous referees for valuable comments.

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Contemporary Economic Policy
(ISSN 1074-3529)
Vol. 23, No. 2, April 2005, 224–241

doi:10.1093/cep/byi017

Additional housework responsibilities could arise from other sources, such as homeownership, which may be associated with specific stages of the life cycle. As the size of the family, the age of the children, and the probability of homeownership change over the life cycle, so, too, will the required amount of housework.

The timing of housework tasks also may change over the life cycle. Some household chores must be performed continuously during the week, for example, meal preparation; whereas others can wait until the weekend, for example, yard work. Individuals who can postpone their housework tasks to nonwork days may not face the same energy constraint as their counterparts who are required to perform housework on workdays. Housework time may have differential wage effects, for the same number of hours per week, for individuals who perform it continually throughout the week versus those who regard it as primarily a weekend activity. There may be less flexibility in the timing of housework if children, especially young children, reside in the home. For this reason, there may be life-cycle effects associated with the timing of housework as well.

There is also the possibility that the intensity of effort associated with an hour of housework time changes over the life cycle. The effort intensity associated with housework is likely to be a function of the type of housework performed. Some chores call for more effort than do others, for example, emptying the dishwasher versus cooking the meal; paying the bills versus mowing the lawn, and so on. Furthermore, the effort associated with typical housework tasks (meal preparation, cleaning, etc.) may be more intensive when combined with the presence of young children. As the nest starts to empty, the effort associated with certain types of housework is likely to decrease.

This study examines these potential life-cycle effects by analyzing the relationship between housework time (hours per week) and wages for samples of young, middle-aged, and older married women and men using samples from the Panel Study of Income Dynamics (PSID). The PSID is a panel study that has annually (from 1968 to 1997, biennially thereafter) collected information on a representative sample of U.S. individuals and the family units in which they reside (Institute for Social Research, 2003a). A unique feature of the PSID is that it includes annual and weekly measures of husbands' and wives' housework time.

This article proceeds as follows. In the next section the authors discuss related research, paying particular attention to a 1997 study by Hersch and Stratton. Hersch and Stratton's study is a comprehensive examination of the effect of housework time on the wages of married men and women, and as such, it is a good starting point for this research. The article then presents the empirical model, which is followed with a discussion of the estimation techniques used in the analysis. The third section contains a description of the variables along with a discussion of the data issues dictated by use of the PSID. Following that the authors present the empirical results, a discussion of the primary findings, and a wage decomposition analysis. The final section contains some conclusions.

## II. RELATED RESEARCH

There is no evidence to suggest that there has been a significant change in the division of labor within the household since Becker published his article in 1985. Despite a slight reallocation of housework activities from wives to husbands in recent years, most of the housework as well as the care of children within the home are still primarily the responsibility of the woman (Hochschild, 1989; Walzer, 1998; U.S. Department of Education, 2001). Furthermore, there continues to exist a high degree of specialization in the type of housework tasks performed by husbands and wives (see, for example, Greenstein, 2000). Wives are still primarily doing "women's work," such as meals, dishes, cleaning, shopping, and laundry, whereas men are engaged in the traditional male tasks such as yard work, home maintenance, and auto repair.

Studies that have examined the relationship between housework and wages have reported a significant negative effect of housework time on wages using ordinary least squares (OLS) models. Wage penalties associated with housework have been shown to hold for female piece-rate workers (Hersch, 1985), self-employed women (Hundley, 2000), and samples of men and women from industrialized countries other than the United States (McAllister, 1990). Researchers have

concluded that household labor time is as important as occupational choice in explaining men's higher earnings (McAllister, 1990) and believe that controlling for it can explain a significant portion of the gender wage gap (Hersch and Stratton, 1997, 2002; Shelton and Firestone, 1989).

These studies differ in terms of who they find suffers adverse wage effects from performing housework. Coverman (1983) reports that household production is negatively related to the wages of white husbands and wives. Others, such as Hersch (1991a,b) and Hersch and Stratton (1997, 2002), find consistent evidence that housework time reduces women's but not men's wages. Alternatively, McLennan (2000) finds little evidence of a housework effect on the wages of any group of male or female workers (white, black, single, or married).

This article closely follows Hersch and Stratton's (1997) study on housework and the wages of married men and women. In their study, Hersch and Stratton address many issues including whether housework time and wages are jointly determined, the impact of unobserved heterogeneity on wage/housework effects, and specification issues such as nonlinear wage/housework functions and threshold effects. They conclude that hours spent on housework adversely affect married women's wages, have an indeterminate effect on married men's wages, and increase the explained component of the gender wage gap by 8–12 percentage points. They also present strong evidence that women's wages and housework time are jointly determined.

Following their approach, the authors pay particular attention to the importance of unobserved heterogeneity as well as potential endogeneity issues. They do so because it is possible that some of these issues could vary with age. For example, there may be more variation in the cognitive abilities or health status of older workers than that of young or middle-aged workers. If this is the case, older workers may be a relatively more heterogeneous group in terms of the amount of physical or mental energy each member can expend on home or market work. Furthermore, if the presence of children affects the timing or the type of housework activities, the correlation between housework time and individual-specific characteristics such as ability may be greater during stages of the life cycle that are associated with fertility than it would be at other stages.

The authors use the same panel study as did Hersch and Stratton, the PSID, although they update the sample period to include more recent data. They also expand the empirical analysis to include an instrumental variables procedure for panel data that accounts for both heterogeneity and endogeneity. The next section describes the empirical model focusing on the issues of heterogeneity and endogeneity.

## III. THE EMPIRICAL MODEL

Most studies that have analyzed the direct effect of housework on wages estimate a standard wage equation augmented with a housework-time variable. The theoretical basis for using an augmented wage equation to test for the effect of performing housework on wages derives from Becker's (1985) allocation of effort model, in which he suggests that housework may directly affect an individual's wage by limiting the amount of energy and effort he or she can expend on the job. Because individual effort is limited and must be allocated across all activities, individuals who allocate a relatively greater proportion of their energy toward childcare and housework will have less energy to allocate toward market work. Less energy or effort expended on the job may result in lower pay associated with the lower productivity. Alternatively it could result in the acceptance of a job that pays less because it calls for less productivity. In either case, a wage gap may develop between workers who have few household responsibilities and their more burdened counterparts.

To determine whether housework time adversely affects wages, the authors augment a standard wage equation with a housework time variable. The wage equation becomes:

$$(1) \qquad \ln W_{it} = \beta X_{it} + \gamma Z_i + \delta HW_{it} + e_{it},$$

where $\ln W_{it}$ is the natural log of the real hourly wage of individual $i$ at time $t$; $X_{it}$ is a vector of measurable human capital and job-related characteristics that vary over time; $Z_i$ is a vector of time-invariant individual characteristics such as gender or race; $HW_{it}$ is weekly time spent on household activities of individual $i$ at time $t$; and $e_{it}$ is the error term.

Hersch and Stratton (1997) emphasize that using OLS to estimate the wage equation could result in biased and/or inefficient estimates of the housework coefficient, $\delta$, for reasons of heterogeneity and endogeneity. This would occur if the error term, $e_{it}$, is actually, $e_{it} = \mu_i + \varepsilon_{it}$, where $\mu_i$ is an unobserved individual-specific effect and $\varepsilon_{it}$ is a random error. $\mu_i$ is assumed to account for individual effects, such as ability, that are not included in the regression, and to be time-invariant and independent over the panels. Alternatively, $\varepsilon_{it}$ is assumed to be uncorrelated with $\mu_i$ and the explanatory variables and to vary independently across individuals and over time (see, for example, Baltagi, 1995).

The heterogeneity issue exists because the individual effect, $\mu_i$, could vary systematically across individuals. In this case, assuming $\mu_i$ is not correlated with any of the explanatory variables, OLS would produce consistent estimates of the parameters, but these estimates would be inefficient compared to those from a procedure that takes this heterogeneity into account. Alternatively, biased estimates of $\delta$ would occur if $\mu_i$ and housework time are correlated.[1] Assume that $\mu_i$ measures innate ability and individuals with greater values of $\mu_i$ receive higher wages, perhaps because they are more productive in the market. If these individuals also spend less time on housework, then $\mu_i$ and housework time will be negatively correlated. In this case, the estimated housework coefficient from a cross-sectional OLS wage equation will suffer from "heterogeneity" bias, and time spent on housework will appear to have a greater negative effect on wages than it actually does.[2]

The authors investigate these issues by first using OLS to obtain benchmark estimates of $\delta$, and then compare those estimates to ones obtained from procedures used to control for heterogeneity and endogeneity. The first procedure used is the fixed effects (within) estimator, which controls for the potential endogeneity by effectively removing the individual effect, $\mu_i$, from the regression model. Assuming the random error, $\varepsilon_{it}$, is homoscedastic and uncorrelated with any of the explanatory variables, the fixed effects procedure results in unbiased estimates of the time-varying variables. Its major drawback is that it also removes all time-invariant variables, such as gender and race, so their coefficients cannot be estimated. Furthermore, under certain circumstances, efficiency can be improved with an instrumental variables (IV) estimator (Baltagi, 1995).

The second procedure used is the Hausman-Taylor instrumental variables (HTIV) estimator (Hausman and Taylor, 1981), which assumes that a subset of the explanatory variables is correlated with the individual effect, but the random error remains uncorrelated with the explanatory variables. Because panel data consist of pooled observations on the same individuals over different periods, instruments can be derived from within the model (Hausman and Taylor, 1981). The following description of the HTIV estimator comes from the xthtayor procedure in Stata 8 (Stata Corporation, 2003a, pp. 95–107).

Consider the following wage equation:

$$(2)\quad \ln W_{it} = \beta_1 X_{1it} + \beta_2 X_{2it} + \gamma_1 Z_{1i} + \gamma_2 Z_{2i} + \delta HW_{it} + \mu_i + \varepsilon_{it},$$

for $i = 1, \ldots n$ and for each $i$, $t = 1, \ldots T_i$, of which $T_i$ periods are observed and $n$ is the number of individuals in the sample. In equation (2), $X_{1it}$ includes all the time-varying *exogenous* variables; $X_{2it}$ includes the time-varying *endogenous* variables with the exception of housework time; $Z_{1i}$ includes all the time-invariant *exogenous* variables; and $Z_{2i}$ includes all the time-invariant *endogenous* variables. All else is defined as before.

The HTIV procedure is implemented via a standard generalized least squares (GLS) transformation of the dependent and independent variables using a weight constructed from estimates of the variances of $\mu_i$ and $\varepsilon_{it}$. The estimates of the variances are obtained from within estimates of the $\beta$s and intermediate IV estimates of the $\gamma$s. Next, an IV regression is fitted on the transformed GLS variables using the following as instruments: the within-person means of the time-varying exogenous variables ($\bar{X}_{1i} = \Sigma_{t=1}^{Ti} X_{1it}/n$),

1. Others have addressed the potential correlation between $e_{it}$ and housework time. See, for example, Hersch and Stratton (1997, 2002) and McLennan (2000).

2. Heterogeneity bias is the bias caused by omitting a variable that is correlated with the explanatory variable (e.g., Wooldridge, 2003).

deviations from the within-person means of the time-varying exogenous and endogenous variables ($X_{1i}t - \bar{X}_{1i}$ *and* $X_{2it} - \bar{X}_{2i}$), and the time-invariant exogenous variables ($Z_{1i}$). As long as $\varepsilon_{it}$ is not correlated with the explanatory variables, deviations from the mean are valid instruments because they are uncorrelated with the error term by construction. If the equation is overidentified, that is, if the number of time-varying exogenous variables ($X_{1it}$) exceeds the number of time-invariant endogenous variables ($Z_{2i}$), then the HTIV estimator is consistent and more efficient than the fixed effects estimator (Baltagi, 1995).

## IV. THE DATA

### A. *The Variables*

As stated in the introduction, the analysis is based on samples drawn from the PSID. Hersch and Stratton's sample consists of white, married, 20–64-year-olds, and their sample period is restricted to the 1979–87 waves of the PSID. Similarly, the authors restricted the sample to married individuals who are 20 to 65 years old but drew them from the 1983–93 waves. The authors also expanded the sample to include nonwhite individuals.

A unique feature of the PSID is that it provides an annual measure of the weekly housework time of husbands and wives.[3] This information is collected from respondents' answers to the following questions: "About how much time do you spend on housework in an average week? I mean time spent cooking, cleaning, and doing other work around the house" and "About how much time does your wife spend on housework in an average week? I mean time spent cooking, cleaning, and doing other work around the house" (Institute for Social Research, 2003b). Note that this housework variable does not include child care, and as such, the authors believe it understates the actual home production time of individuals with children.

The hourly wage variable used in this article is constructed from heads' answers to questions about their wages and their wives' wages at their *current* jobs.[4] The authors converted all wage and income variables to real wages and income using the monthly Consumer Price Index for all urban consumers (1982–84 = 100) and imposed a lower limit of $1.00 per hour on the hourly wage variable.

The authors add a measure of market time in the empirical model because they believe doing so provides a better test of Becker's hypothesis that those with heavier household responsibilities will have less energy available for the market. They follow McLennan (2000), who argues the true test of Becker's model calls for directly controlling for the number of market hours in the wage equation.[5] In Becker's model, holding the number of market hours constant, an increase in the number of hours spent in home production will reduce the effort per hours spent in the market, which should negatively affect wages. Housework time can be used as a proxy for market effort only if one controls for the number of market hours. Otherwise, individuals could respond to an increase in their housework time by reducing their market hours and the amount of effort expended per hours on the job would not change. In that case, there would be no correlation between market effort and housework time.

Unfortunately, with the exception of the 1985 through 1987 waves, the number of hours worked per week at the respondent's main job refers to the number of hours worked per week in the previous year, rather than the current year. For this reason, the market-time variable is a proxy for current market time

3. According to the PSID, information regarding each family member is obtained through answers provided by the head of the family unit, with the exception of selected survey waves (e.g., 1976, 1985) when the PSID administrated separate spousal surveys. Initially, to be consistent with Census definitions at the time, the husband was designated as the head of the family. In recent surveys, the male is still likely to be the head unless he is unable to fulfill the functions of head or his female counterpart insists on being the head (Institute for Social Research, 2003a). Therefore, it is likely that most of the information for the sample of wives comes from their husbands, although it is possible that in some cases the wife could be designated as the head and provide information about her husband to the PSID.

4. It is not the average hourly earnings variable constructed by the PSID staff, which is based on total labor income and the total number of hours of market time from the *previous* year and, as such, is not contemporaneously related to the housework variable.

5. Hersch and Stratton indirectly control for market hours by imposing a sample restriction of full-time employment and find similar results for the restricted and unrestricted samples.

and is based on the number of hours per week spent, on average, at the main job in the previous year.[6] For individuals for whom that information is missing, the authors define market hours as the average number of hours per week worked at all jobs in the previous year.

Other variables used in the analysis include the standard human capital variables (education, experience, and tenure), job-related variables, regional variables, and variables to control for economy-wide fluctuations. Although most of these variables are measured in the standard way, a few need some clarification. Because the PSID data do not contain actual experience for each year, the authors computed the experience variable by augmenting 1976 work experience with the number of hours worked in each successive year divided by 2,000.[7] There are two things to note about this measure of experience. First, because the number of annual hours worked is based on hours worked in the previous year, it does not include any experience gained during the current year. Second, if an individual worked more than 2,000 hours in a given year, he or she would have more than a year's worth of experience in that year. For these reasons, average years of experience in this study may differ from averages presented by other studies using the PSID.

The authors include a disability variable equal to one if the individual has a physical or a nervous condition that would limit the type or the amount of (home and market) work he or she could do. It is zero otherwise. To control for urban and rural wage differentials, the authors include a big city variable that is one if the individual resides in a county where the largest city has a population of 500,000 or more, and zero otherwise. Finally, they include a union variable that is based on whether the individual's current job is covered by a union contract rather than whether he or she is a member of a union.

### *B. The Samples of Young, Middle-Aged, and Older Men and Women*

Because the authors are interested in examining whether the effect of housework time on wages varies with age, they segment the sample into three age groups: 20 to 34, 35 to 49, and 50 to 65-year-olds. Table 1 presents the means and standard deviations of selected variables for samples of employed husbands and wives separately for each age group. Although the descriptive statistics are based on the number of job-years rather than the number of individuals, the authors adjusted the sampling weights to account for the number of job-years by dividing the sampling weight by the number of times each individual falls into one of the age groups. Controlling for age cohort, a comparison of the male and female sample means reveals typical male/female results. On average, employed women earn less per hour, perform more hours per week of housework, work fewer hours per week in the market, have less prior experience and less current tenure, and are less likely to be covered by a union contract than their male counterparts.

Of particular interest are the differences in hours of housework performed per week by older and younger workers. Young women (20–34) do ~ two fewer hours of housework per week than do middle-aged and older women. Older men (50–65) do one hour less of housework per week than do middle-aged and younger men. Among the men, those 50–65 are the most likely to perform less than 10 hours of housework per week, whereas among the women, those in the youngest age group are the most likely to perform less than 10 hours of housework per week. Furthermore, the female/male ratio of housework time increases with age. Young women spend 2.4 times more hours on housework per week than do young men, whereas older women spend 3.1 times more hours on housework than do older men. All this suggests that an individual's location on his or her wage-housework function may depend on age as well as gender.

To address whether these differences in young and older workers' average housework time and in their housework-time

6. The authors assume that lagged market time is a reasonable control of current market time. To explore this issue, they calculated the correlation between current and lagged market hours for the years both are reported (1985 through 1987). For employed women, the average (sample-weighted) correlation between current market hours and lagged market hours is 0.61; for employed men, it is 0.51.

7. The 1976 wave was the first to include information on wives' prior experience directly through a separate spouse questionnaire. For those who established households after 1976, the authors used the first year for which information on experience was available.

**TABLE 1**
Descriptive Statistics of Selected Variables for Samples of Employed Married Women and Men, Segmented by Age, 1983–93

| | WOMEN | | | | | | MEN | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Age Group** | **20–34** | | **35–49** | | **50–65** | | **20–34** | | **35–49** | | **50–65** | |
| **Variable** | **Mean** | **SD** | **Mean** | **SD** | **Mean** | **SD** | **Mean** | **SD** | **Mean** | **SD** | **Mean** | **SD** |
| Hourly wage (1982–84 dollars) | 7.25 | 10.24 | 8.26 | 11.30 | 7.94 | 16.27 | 9.49 | 9.32 | 12.68 | 9.21 | 13.35 | 13.688 |
| Housework (hours/week) | 17.38 | 11.62 | 19.14 | 11.06 | 18.91 | 10.25 | 7.26 | 6.99 | 7.16 | 7.05 | 6.11 | 6.572 |
| *Percent who performed* | | | | | | | | | | | | |
| 9 or fewer hours of housework/week | 21.9% | 0.41 | 14.7% | 0.36 | 13.5% | 0.34 | 67.2% | 0.47 | 66.7% | 0.47 | 73.1% | 0.44 |
| 10–19 hours of housework/week | 38.0% | 0.49 | 35.4% | 0.48 | 37.4% | 0.48 | 24.5% | 0.43 | 24.5% | 0.43 | 20.4% | 0.40 |
| 20 or more hours of housework/week | 40.1% | 0.49 | 49.9% | 0.50 | 49.1% | 0.50 | 8.3% | 0.28 | 8.8% | 0.28 | 6.4% | 0.25 |
| Age | 28.59 | 4.14 | 40.77 | 4.68 | 54.9 | 4.31 | 29.28 | 3.90 | 40.71 | 4.68 | 55.71 | 4.65 |
| *Race/ethnicity* | | | | | | | | | | | | |
| White | 84.4% | 0.36 | 85.8% | 0.35 | 86.1% | 0.35 | 84.8% | 0.36 | 87.3% | 0.33 | 87.2% | 0.33 |
| Black | 8.6% | 0.28 | 7.7% | 0.27 | 7.8% | 0.27 | 9.1% | 0.29 | 7.8% | 0.27 | 7.5% | 0.26 |
| Other | 7.0% | 0.26 | 6.4% | 0.25 | 6.0% | 0.24 | 6.0% | 0.24 | 4.9% | 0.22 | 5.3% | 0.22 |
| Education (years completed) | 13.46 | 2.07 | 13.33 | 2.28 | 12.62 | 2.37 | 13.31 | 2.23 | 13.54 | 2.54 | 12.82 | 3.05 |
| Market time (lagged hours/week) | 35.93 | 11.37 | 35.04 | 11.65 | 34.01 | 11.70 | 43.65 | 9.72 | 44.37 | 9.61 | 42.66 | 10.17 |
| Experience (years) | 6.92 | 4.19 | 13.23 | 6.95 | 18.78 | 10.75 | 9.98 | 4.53 | 20.96 | 6.13 | 34.97 | 7.80 |
| Tenure (years) | 3.67 | 3.61 | 6.70 | 6.29 | 10.10 | 8.18 | 4.63 | 4.17 | 9.31 | 7.89 | 15.10 | 12.61 |
| Union contract | 13.7% | 0.34 | 15.7% | 0.36 | 15.2% | 0.36 | 20.8% | 0.41 | 24.7% | 0.43 | 27.9% | 0.45 |
| Big city (population $\geq$ 500,000) | 16.0% | 0.37 | 14.5% | 0.35 | 13.9% | 0.35 | 15.6% | 0.36 | 15.8% | 0.36 | 14.9% | 0.36 |
| Health limitation | 6.2% | 0.24 | 9.0% | 0.29 | 13.0% | 0.34 | 5.9% | 0.24 | 9.3% | 0.29 | 14.6% | 0.35 |
| *Geographic region* | | | | | | | | | | | | |
| Northeast | 21.2% | 0.41 | 22.5% | 0.42 | 31.3% | 0.46 | 20.7% | 0.41 | 22.6% | 0.42 | 26.6% | 0.44 |
| North central | 26.4% | 0.44 | 27.1% | 0.44 | 26.9% | 0.44 | 26.1% | 0.44 | 27.2% | 0.45 | 27.2% | 0.45 |
| South | 31.6% | 0.47 | 30.8% | 0.46 | 24.9% | 0.43 | 31.8% | 0.47 | 29.5% | 0.46 | 28.5% | 0.45 |
| West | 19.9% | 0.40 | 18.8% | 0.39 | 16.8% | 0.37 | 20.2% | 0.40 | 19.6% | 0.40 | 17.1% | 0.38 |
| *Number of children* | | | | | | | | | | | | |
| Total number (1–17 years old) | 1.19 | 1.18 | 1.31 | 1.18 | 0.19 | 0.56 | 1.30 | 1.16 | 1.56 | 1.20 | 0.34 | 0.73 |
| Younger than 3 | 0.29 | 0.51 | 0.08 | 0.29 | 0.01 | 0.13 | 0.40 | 0.58 | 0.15 | 0.39 | 0.02 | 0.15 |
| 3–5 years old | 0.27 | 0.50 | 0.12 | 0.36 | 0.01 | 0.13 | 0.34 | 0.56 | 0.20 | 0.44 | 0.02 | 0.15 |
| 6–13 years old | 0.54 | 0.87 | 0.66 | 0.87 | 0.05 | 0.25 | 0.50 | 0.81 | 0.81 | 0.93 | 0.12 | 0.40 |
| 14–17 years old | 0.09 | 0.33 | 0.45 | 0.65 | 0.12 | 0.34 | 0.06 | 0.26 | 0.40 | 0.64 | 0.18 | 0.44 |
| Number of job-years | 12,452 | | 10,103 | | 3,209 | | 14,099 | | 13,057 | | 5,128 | |
| Number of individuals | 3,697 | | 2,620 | | 831 | | 3,870 | | 3,108 | | 1,266 | |

*Notes:* A lower limit of $1.00 per hour in 1982–84 dollars is imposed on the hourly wage variable. The means are sample-weighted, with the weights adjusted to account for the number of job years.

*Source:* PSID 1983–93.

**TABLE 2**
OLS Determinants of the Hourly (Log) Wage of Married Women and Men, 1983–93

| | Women | | | Men | | |
|---|---|---|---|---|---|---|
| **Age Group** | **20–34** | **35–49** | **50–65** | **20–34** | **35–49** | **50–65** |
| Housework (hours/week) | −0.411*** | −0.373*** | −0.412*** | −0.169*** | −0.193** | −0.389*** |
| | (0.047) | (0.060) | (0.092) | (0.055) | (0.084) | (0.139) |
| *Race/ethnicity* | | | | | | |
| White | — | — | — | — | — | — |
| Black | −11.448*** | −12.293*** | −19.358*** | −17.620*** | −17.824*** | −15.988*** |
| | (1.397) | (2.139) | (3.914) | (1.515) | (1.991) | (3.553) |
| Other | 0.627 | −0.417 | −2.093 | −6.727** | −5.392* | −2.464 |
| | (2.353) | (2.872) | (4.973) | (2.907) | (3.148) | (5.084) |
| Education (years) | 4.623*** | 1.901*** | 2.008** | 3.776*** | 3.043*** | 1.897** |
| | (0.506) | (0.707) | (0.967) | (0.437) | (0.622) | (0.842) |
| Market time (lagged hours/week) | 0.339 | 0.163 | 0.129 | 1.300*** | 1.400*** | 1.044*** |
| | (0.250) | (0.310) | (0.482) | (0.198) | (0.336) | (0.392) |
| Market time squared | −0.002 | 0.002 | 0.002 | −0.013*** | −0.016*** | −0.011** |
| | (0.004) | (0.005) | (0.007) | (0.002) | (0.004) | (0.005) |
| Experience (years) | 2.151*** | 1.205*** | 0.567 | 1.892*** | 0.454 | 1.546** |
| | (0.491) | (0.372) | (0.451) | (0.555) | (0.603) | (0.621) |
| Experience squared | −0.023 | −0.007 | −0.006 | −0.062** | −0.003 | −0.021* |
| | (0.031) | (0.012) | (0.010) | (0.026) | (0.016) | (0.012) |
| Tenure (years) | 5.199*** | 3.467*** | 2.110*** | 3.363*** | 1.096*** | 0.860** |
| | (0.420) | (0.345) | (0.390) | (0.365) | (0.325) | (0.348) |
| Tenure squared | −0.228*** | −0.072*** | −0.019 | −0.142*** | −0.006 | 0.003 |
| | (0.033) | (0.016) | (0.013) | (0.026) | (0.013) | (0.009) |
| Union contract | 18.125*** | 10.310*** | 14.965*** | 20.326*** | 9.890*** | 10.445*** |
| | (1.527) | (1.818) | (3.097) | (1.278) | (1.667) | (2.533) |
| *Region of residence* | | | | | | |
| Northeast | — | — | — | — | — | — |
| North central | −12.793*** | −15.838*** | −17.155*** | −8.703*** | −11.936*** | −11.622** |
| | (2.191) | (2.710) | (3.817) | (2.289) | (2.747) | (4.738) |
| South | −10.923*** | −13.300*** | −18.530*** | −9.751*** | −12.087*** | −12.370*** |
| | (1.893) | (2.545) | (3.875) | (1.998) | (2.528) | (4.428) |
| West | −1.142 | −3.612 | −5.037 | −0.764 | −5.138* | −2.265 |
| | (2.133) | (2.808) | (4.341) | (2.305) | (2.861) | (4.719) |
| Big city (≥ 500,000) | 10.850*** | 10.782*** | 13.416*** | 6.089*** | 11.520*** | 12.175*** |
| | (1.467) | (1.863) | (3.327) | (1.475) | (1.978) | (3.641) |
| Health limitation | −3.213* | −1.922 | −0.414 | −10.108*** | −13.985*** | −11.486*** |
| | (1.879) | (2.226) | (3.076) | (1.912) | (2.169) | (2.651) |
| Constant | 81.406*** | 32.048 | 299.555 | 90.275** | −69.161 | −107.261 |
| | (29.476) | (52.258) | (183.962) | (35.091) | (54.503) | (140.106) |
| Number of job years | 12,452 | 10,103 | 3,209 | 14,099 | 13,057 | 5,128 |
| Number of individuals | 3,697 | 2,620 | 831 | 3,870 | 3,108 | 1,266 |
| Adjusted $r^2$ | 0.3837 | 0.4247 | 0.3924 | 0.3581 | 0.3882 | 0.4247 |

*Notes:* The estimates and robust standard errors (in parentheses) are multiplied by 100. The regressions also include age; age squared; the unemployment rate by region of residence; the average wage by year, age, gender, and level of education; the number of years of education of the individual's mother, father, and spouse; and a set of year dummy variables (1983 to 1993). ***, **, * Significant at the 0.01, 0.05, and 0.10 levels, respectively.

distributions justify segmenting the wage equations by age, the authors used Chow tests to test for age-based structural differences in housework time. The Chow tests are based on OLS regression models pooled across age, using the same explanatory variables used in the segmented OLS models (see Table 2). For the pooled regressions, the authors stratified the housework-time variable into three variables by interacting it with three age-cohort dummy variables. Results from these tests indicate, at the 5% level, that the structure of housework time does vary for these three age groups, controlling for gender.[8] Based on this evidence, it seems reasonable to proceed using the segmented samples.

## V. THE EMPIRICAL RESULTS

The authors first examine baseline estimates from OLS regressions and then present estimates from fixed effects and HTIV models. A comparison of the housework-time estimates from the three models along with the results from Hausman specification tests (Hausman, 1978) allows the authors to address the issues of unobserved heterogeneity and endogeneity. Finally, they include a wage decomposition analysis, which is performed separately for the three age groups. Based on the OLS and IV estimates, the authors calculate how much (if any) of the male/female wage gap can be explained by gender differences in housework time. All models were estimated using Stata 8 (Stata Corporation, 2003b).

Because there are multiple observations on individuals, the within-person errors are likely to be correlated. For this reason, the authors corrected the standard errors in the OLS regressions using the Huber-White estimator of variance with a correction for within-person correlation (Rogers, 1993; White, 1980).[9] In general, the corrected standard errors are larger than the uncorrected standard errors, which suggest *t*-statistics calculated from the uncorrected standard errors may overstate the effect of the independent variables on wages. Finally, it is important to note that this analysis is not able to distinguish life-cycle effects from cohort effects, and as such, some differences attributed to life-cycle effects may have been influenced by changes over time in societal expectations and/or norms. The authors elaborate on this point in the conclusions.

### *A The OLS Results*

Table 2 includes the OLS estimates of equation (1) for the samples segmented by age and gender. In addition to the standard wage determinants, the analysis included a few additional variables that were used as instruments in the HTIV equations. These variables are age; age squared; years of education of the individual's mother, father, and spouse; a regional dummy variable (Northeast, North central, South, and West); and an average wage based on an annual Current Population Survey earnings series of median annual income presented separately by gender, age, and years of school completed. Columns 1 through 3 contain these estimates for the three samples of women, and columns 4 through 6 include these estimates for the men's samples.[10] For most samples, the standard wage determinants have the expected signs and levels of significance. Years of education, experience and tenure, union status, and residing in or near a large city are positively related to wages. Regardless of gender or age, being black and residing in the North central or Southern regions of the United States result in lower wages. For men, lagged weekly hours of market time tend to increase wages at a decreasing rate, whereas for women, the lagged hours of market time estimates are not significantly related to wages at the 5% level. With

8. The *F*-statistics constructed from the pooled and separate regressions are, for the female sample, $F(65, 7{,}046) = 1.75$, and for the male sample, $F(65, 8{,}142) = 1.73$, where the critical value, at the 5% level, is 1.34. Estimates from these models are available from the authors.

9. The Huber-White estimator of variance does not correct for the original clustering and stratification of the sample. To the extent there is within-cluster correlation (i.e., individuals in each cluster are statistically similar to their neighbors), the reported standard errors may be underestimated. The authors are grateful to Daniel H. Hill for this point.

10. The authors did examine the extent of sample selection bias in the OLS estimates by estimating selection-corrected models (Heckman, 1979). With the exception of the older women sample, the estimate of lambda was positive and significant; however, an examination of the estimated coefficients shows little evidence that sample selection bias affected the estimates of any of the independent variables especially the housework estimates. For this reason, the fixed effects and HTIV models do not take potential sample selectivity into account. (These estimates are available from the authors.)

a few exceptions, these results hold for the other wage analyses.

For all groups, the analysis finds that an additional hour of housework reduces wages by a statistically significant but small amount. For women, performing an additional hour of housework per week reduces their wages by ~ 0.4%. For men, performing an additional hour of housework reduces their wages by 0.1–0.4%. These results are similar to those reported by other studies in that OLS models yield small but statistically significant housework-time effects.

### *B. The Fixed Effects Results*

Table 3 includes the estimates from the fixed effects regressions. As in the case of the OLS estimates, the authors find that most of the standard wage determinants have the expected signs and levels of significance,

**TABLE 3**
Fixed Effects Determinants of the Hourly (Log) Wage of Married Women and Men, 1983 to 1993

| | Women | | | Men | | |
|---|---|---|---|---|---|---|
| **Age Group** | **20–34** | **35–49** | **50–65** | **20–34** | **35–49** | **50–65** |
| Housework (hours/week) | −0.140*** | −0.072** | −0.107 | 0.015 | −0.012 | 0.090 |
| | (0.032) | (0.036) | (0.067) | (0.038) | (0.037) | (0.070) |
| Market time (lagged hours/week) | 0.421*** | 0.270** | −0.553** | 0.669*** | 0.693*** | 0.415*** |
| | (0.110) | (0.117) | (0.237) | (0.114) | (0.108) | (0.161) |
| Market time squared | −0.006*** | −0.004** | 0.005 | −0.008*** | −0.010*** | −0.006*** |
| | (0.002) | (0.002) | (0.004) | (0.001) | (0.001) | (0.002) |
| Experience (years) | 1.912*** | 1.450*** | −0.016 | 3.556*** | 1.303*** | 1.203*** |
| | (0.438) | (0.338) | (0.433) | (0.387) | (0.344) | (0.458) |
| Experience squared | −0.064*** | −0.027*** | 0.007 | −0.111*** | −0.011 | −0.016** |
| | (0.022) | (0.010) | (0.009) | (0.015) | (0.008) | (0.008) |
| Tenure (years) | 2.792*** | 1.990*** | 1.128*** | 1.681*** | 1.539*** | 0.862*** |
| | (0.281) | (0.204) | (0.298) | (0.222) | (0.143) | (0.179) |
| Tenure squared | −0.143*** | −0.054*** | −0.012 | −0.112*** | −0.041*** | −0.003 |
| | (0.022) | (0.009) | (0.010) | (0.017) | (0.006) | (0.005) |
| Union contract | 9.631*** | 3.929*** | 11.641*** | 11.732*** | 6.435*** | 9.252*** |
| | (1.314) | (1.368) | (2.803) | (1.033) | (1.013) | (1.901) |
| *Region of residence* | | | | | | |
| Northeast | — | — | — | — | — | — |
| North central | −1.272 | −12.159** | −28.650* | 4.785 | −8.241** | −0.362 |
| | (4.194) | (5.789) | (15.539) | (3.170) | (3.210) | (7.265) |
| South | −6.657** | −14.992*** | −39.724 | −1.057 | −5.453* | −10.795 |
| | (3.390) | (5.190) | (26.002) | (2.397) | (2.792) | (8.050) |
| West | 9.431** | −10.577* | −4.479 | 0.436 | −7.383** | −6.004 |
| | (4.149) | (5.569) | (0.18597) | (2.872) | (3.171) | (8.298) |
| Big city (≥ 500,000) | −5.605** | 9.046*** | −7.357 | −1.257 | 3.614** | 3.900 |
| | (2.312) | (2.533) | (9.355) | (1.759) | (1.815) | (4.356) |
| Health limitation | −3.289** | −2.307* | 1.999 | −2.054 | −1.096 | −1.146 |
| | (1.387) | (1.281) | (1.957) | (1.303) | (1.000) | (1.450) |
| Constant | 150.306*** | 168.985*** | 199.472*** | 157.247*** | 195.403*** | 191.879*** |
| | (3.730) | (5.097) | (0.15287) | (3.772) | (5.098) | (9.400) |
| Number of job years | 12,452 | 10,103 | 3,209 | 14,099 | 13,057 | 5,128 |
| Number of individuals | 3,697 | 2,620 | 831 | 3,870 | 3,108 | 1,266 |
| $r^2$ | 0.0843 | 0.0576 | 0.0451. | 0.1120 | 0.0558 | 0.0652 |
| Exogeneity test statistic | $\chi^2_{22} = 301$ | $\chi^2_{22} = 87$ | $\chi^2_{22} = 67$ | $\chi^2_{22} = 100$ | $\chi^2_{22} = 284$ | $\chi^2_{22} = 118$ |

*Notes:* The estimates and standard errors (in parentheses) are multiplied by 100. The regressions also include a set of year dummy variables (1983 to 1993). The exogeneity test examines whether the fixed effect is correlated with some of the explanatory variables under the null assumption there is no correlation. The critical value of the test statistic is 33.92 at the 5% level. ***, **, * Significant at the 0.01, 0.05, and 0.10 levels, respectively.

although the point estimates are smaller in magnitude than that of their OLS counterparts. One change is the positive and significant relationship between young and middle-aged women's (lagged) market time and their current wages.

An examination of the housework-time estimates shows the effect of housework time on wages is significant for women who are younger than 50. Once the authors control for the individual effect, young and middle-aged women appear to be the only ones who incur wage losses from performing housework. Their losses are small but statistically significant. Each additional hour of housework time reduces their wages by 0.14–0.07%, an amount that is one-third to one-fifth the size of their OLS estimates. As in the case of the OLS estimates, the fixed effects estimates are very similar to those presented by Hersch and Stratton. Although the OLS estimates indicated a negative housework effect on wages for men and older women, their fixed effects housework estimates are insignificant. Once the individual effect is controlled, housework time appears to have no independent effect on wages.

To determine whether one can improve on these estimates with an IV procedure, the authors first implement Hausman's specification test to determine if some of the explanatory variables are correlated with the fixed effect. In this test, the authors compare the fixed effects estimator, which is assumed to be consistent under the null assumption of exogeneity, to a GLS estimator, which is efficient only under the assumption of exogeneity (e.g., Baltagi, 1995). This test generates a chi-square statistic with degrees of freedom equal to the number of coefficients estimated by both models. If there exists a systematic difference between the two estimators, the authors cannot accept the assumption of exogeneity between $\mu_i$ and all of the explanatory variables, including housework time. In this case, the chi-square test statistic will be greater than its critical value.

To implement this test, the authors compared the estimates from the fixed effects procedure presented in Table 3 to those from a random effects procedure that included all of the variables used to estimate the OLS model (see Table 2). The test statistic, which is a chi-square statistic with 36 degrees of freedom under the null hypothesis, has a critical value of 50.99 at the 5% level. Based on the results from this test, which are presented in the last row of Table 3, the authors cannot accept the hypothesis of exogeneity between the individual effect and all of the explanatory variables for any age group of men or women. Therefore, it seems reasonable to proceed with the HTIV estimator to see if the authors can improve on the fixed effects results.

### C. *The HTIV Regressions*

The primary challenge of using any IV procedure in an analysis that compares the estimates from multiple samples is coming up with a set of instruments that is not correlated with the individual effect for *all* of the samples. Fortunately, Hausman and Taylor (1981) provide the following specification test, which allows one to determine if the sets of variables identified as exogenous and endogenous are appropriate. If the equation is overidentified, the assumption that certain variables are not correlated with the individual effect can be tested. This test of exogeneity restrictions is a variant of a Hausman test and is based on the difference between the fixed effects and the HTIV estimators. It produces a test statistic that is distributed as a chi-square with degrees of freedom equal to the number of overidentifying restrictions (the number of time-varying exogenous variables minus the number of time-invariant endogenous variables). If there exists a systematic difference between the fixed effects and HTIV estimators, the authors cannot accept the assumption that the variables assumed to be exogenous are all valid instruments. In this case, the chi-square test statistic will be greater than its critical value.

To estimate equation (2), the authors used all the variables included in the OLS model. The set of time-invariant exogenous variables includes two dummy variables for the individual's race (white is the excluded category), mother's education, father's education, and spouse's education. The time-invariant endogenous variable set includes only the individual's years of education. The authors experimented with alternative sets of time-varying exogenous and endogenous variables. For each specification tried, they used the

Hausman test to determine if the set of exogenous variables was a legitimate set of instruments for the assumed set of endogenous variables.

In addition to housework time, the final set of time-varying endogenous variables includes tenure, tenure-squared and union status.[11] Based on the results from the exogeneity restrictions tests, one could consider all other time-varying variables to be a legitimate set of instruments for the samples of men and women with one exception. For the sample of middle-aged men only, to accept the hypothesis that the remaining exogenous variables were legitimate instruments the authors had to exclude age and age squared from the instrument set. The test statistic is a chi-square under the null hypothesis with 10 degrees of freedom for the sample of middle-aged men and 12 degrees of freedom for the other samples. The critical values are 18.3 and 21.0 at a 5% level, respectively. Each group's test statistic is included in the last row of Table 4.

An examination of the HTIV estimates shows the following. In general, relative to the OLS estimates, these results show a smaller impact of race on wages, and a larger return to years of education. Comparing these estimates to those from the fixed effects model generally shows a higher return to market time and, for women, a slight increase in the returns to experience. The returns to tenure, working under a union contract, and having a health limitation (for women only) are similar to the fixed effects estimates.

With respect to the return to housework time, the HTIV estimates continue to show a significant adverse effect of housework time on the wages of young and middle-aged women, and a slight increase in the point estimates relative to the fixed effects estimates. For older women, the housework time estimate becomes marginally significant relative to the fixed effects estimate. Again, for men there is no significant effect of housework time on their wages.

11. Adding in other variables such as experience, experience squared, market time, and market time squared to the list of endogenous variables did not affect the estimates of the housework time variable, although the authors were not able to reject the hypothesis that the remaining variables were legitimate instruments for all six samples.

### *D. Housework and Fertility*

The authors explore one final issue for these models. As stated before, because the housework-time variable does not include the number of hours directly spent on child care, the degree to which it is an accurate measure of the time and effort that an individual spends on home production will depend on that individual's fertility. Excluding measures of fertility from the wage equations could potentially bias the housework-time estimates if children have an exogenous effect on market wages and if children and housework are positively correlated. Many housework chores are complementary to the presence of children without being considered per se child care. For instance, the presence of children will increase the time devoted to laundry, cooking, clean up, grocery shopping, errands, and so on. Because younger couples are more likely to have children in the home than are older couples, the housework-time variable for young individuals may understate both the time and effort associated with actual home production.

To explore this issue, the authors estimated additional OLS, fixed effects, and HTIV regressions that included four variables to control for the number of children in the home. Because the effort associated with childcare may vary with the age of the child, the authors segmented the total number of children reported living in the home into variables that correspond to the following age groups: younger than 3 years, 3–5 years, 6–13 years, and 14–17 years. They estimated two additional specifications of the wage equation for each statistical procedure. The first included the fertility variables but excluded housework time. The second included both housework time and the fertility variables.

For men and older women, the inclusion of the fertility variables in the wage equation did not affect their housework-time estimates, and the inclusion of housework time did not affect their fertility estimates. For young and middle-aged women, including housework time in a wage equation that already includes the fertility variables did reduce the significance of the some of the fertility estimates. The opposite is not true, however. Based on these estimates, it appears that housework time does have a small but independent effect on the wages

**TABLE 4**
HTIV Determinants of the Hourly (Log) Wage of Married Women and Men, 1983–93

| | Women | | | Men | | |
|---|---|---|---|---|---|---|
| **Age Group** | **20–34** | **35–49** | **50–65** | **20–34** | **35–49** | **50–65** |
| Housework (hours/week)† | −0.158*** | −0.093*** | −0.126* | 0.017 | −0.015 | 0.083 |
| | (0.030) | (0.034) | (0.065) | (0.036) | (0.035) | (0.066) |
| *Race/ethnicity* | | | | | | |
| White | — | — | — | — | — | — |
| Black | −8.235*** | −8.446*** | −17.440*** | −16.458*** | −19.578*** | −12.562*** |
| | (1.802) | (2.599) | (5.329) | (1.824) | (2.507) | (4.684) |
| Other | 3.073 | 0.885 | −0.154 | −3.703 | −3.211 | 1.570 |
| | (2.684) | (3.655) | (7.007) | (2.890) | (3.862) | (6.499) |
| Education (years)† | 8.100*** | 7.660*** | 9.628*** | 6.631*** | 6.801*** | 7.788*** |
| | (0.658) | (1.087) | (1.290) | (0.560) | (0.739) | (1.015) |
| Market time (lagged hours/week) | 0.510*** | 0.312*** | −0.364* | 0.763*** | 0.740*** | 0.434*** |
| | (0.099) | (0.107) | (0.218) | (0.104) | (0.101) | (0.150) |
| Market time squared | −0.006*** | −0.004** | 0.004 | −0.009*** | −0.010*** | −0.006*** |
| | (0.001) | (0.002) | (0.003) | (0.001) | (0.001) | (0.002) |
| Experience (years) | 2.718*** | 1.778*** | 0.309 | 3.675*** | 1.183*** | 1.119*** |
| | (0.368) | (0.265) | (0.341) | (0.372) | (0.289) | (0.386) |
| Experience squared | −0.059*** | −0.024*** | 0.003 | −0.112*** | −0.008 | −0.014** |
| | (0.021) | (0.008) | (0.008) | (0.016) | (0.007) | (0.007) |
| Tenure (years)† | 2.968*** | 2.100*** | 1.134*** | 1.788*** | 1.600*** | 0.900*** |
| | (0.265) | (0.194) | (0.288) | (0.210) | (0.137) | (0.170) |
| Tenure squared† | −0.146*** | −0.056*** | −0.010 | −0.115*** | −0.042*** | −0.004 |
| | (0.021) | (0.009) | (0.010) | (0.016) | (0.006) | (0.005) |
| Union contract† | 9.797*** | 4.234*** | 11.798*** | 12.061*** | 6.659*** | 9.753*** |
| | (1.243) | (1.306) | (2.710) | (0.974) | (0.963) | (1.805) |
| *Region of residence* | | | | | | |
| Northeast | — | — | — | — | — | — |
| North central | −9.558*** | −13.802*** | −19.564*** | −2.252 | −9.603*** | −7.379* |
| | (2.146) | (2.810) | (5.465) | (2.010) | (2.261) | (4.247) |
| South | −10.599*** | −14.352*** | −19.144*** | −6.239*** | −9.891*** | −12.318*** |
| | (1.860) | (2.561) | (5.292) | (1.653) | (2.004) | (4.088) |
| West | 2.544 | −3.783 | −3.490 | −0.346 | −6.252*** | −3.330 |
| | (2.159) | (2.869) | (5.861) | (1.917) | (2.254) | (4.595) |
| Big city (≥ 500,000) | 4.788*** | 9.550*** | 9.181** | 2.219* | 6.508*** | 7.709** |
| | (1.434) | (1.781) | (4.416) | (1.267) | (1.435) | (3.024) |
| Health limitation | −3.451*** | −2.499** | 1.512 | −3.281*** | −2.176** | −2.384* |
| | (1.239) | (1.176) | (1.832) | (1.175) | (0.929) | (1.341) |
| Constant | 2.408 | −7.862 | 113.682 | 48.123** | 58.473*** | −151.596* |
| | (24.551) | (34.787) | (119.080) | (23.856) | (8.260) | (79.188) |
| Number of job years | 12,452 | 10,103 | 3,209 | 14,099 | 13,057 | 5,128 |
| Number of individuals | 3,697 | 2,620 | 831 | 3,870 | 3,108 | 1,266 |
| Exogeneity test statistic | $\chi^2_{12} = 8.2$ | $\chi^2_{12} = 19.7$ | $\chi^2_{12} = 12.3$ | $\chi^2_{12} = 1.6$ | $\chi^2_{12} = 9.8$ | $\chi^2_{12} = 10.5$ |

*Notes:* See Table 2. The exogeneity test compares the HTIV estimators to the fixed effects estimators under the null assumption that they do not systematically differ. The critical values of the test statistics are 18.3 for 10 degrees of freedom and 21.0 for 12 degrees of freedom, both at the 5% level. †Endogenous variable. ***, **, * Significant at the 0.01, 0.05, and 0.10 levels, respectively.

of young and middle-aged women.[12] (Appendix Table A-1 includes the housework-time and number-of-children estimates for the young and middle-aged women samples.)

## VI. HOUSEWORK TIME AND THE GENDER WAGE GAP

A few researchers have calculated how much of the male/female wage gap is due to average differences in men's and women's housework time. For example, Shelton and Firestone (1989) report for a cross-sectional sample of husbands and wives from the early 1980s that home production time explains 8.2% of the gender-wage gap. Their measure of home production was derived from time diaries and included time spent on household tasks as well as child care. Hersch and Stratton (1997) report for the period 1979–87 that accounting for husband/wife differences in housework time (minus child care) increases the explained portion of the wage gap by 8–12 percentage points. When they expanded the analysis to a later period, included single workers, and used an expanded measure of housework, they find housework time increases the explained portion of the wage gap by 14 percentage points (Hersch and Stratton, 2002). Based on this evidence, differences in the amount of housework performed by men and women appear to play an important role in understanding men's higher average wages.

To see whether the individual effect and its likely correlation with housework time affects the explained portion of the wage gap, the authors compare the results from two estimators—OLS, which ignores individual effects, and the HTIV procedure, which controls for the individual effect. If the individual effect is correlated with housework time, not accounting for it would indicate that gender differences in housework time explain more of the wage gap than they actually do.

The authors use a variant of the well-known Oaxaca/Blinder wage decomposition to calculate the explained portion of the wage gap (Blinder, 1973; Oaxaca, 1973). Rather than using two sets of estimates from separate male/female wage regressions, they follow Neumark (1988) and use estimates of the coefficients from regressions that are pooled by gender but still segmented by age. The authors estimate two models for each procedure for each age group. The first excludes housework time but includes all the other wage determinants. The second augments the first model by adding in the housework-time variable. Table 5 contains the percentage of the hourly wage difference that is explained by male/female differences in average characteristics.

Most of the results presented in Table 5 are not surprising. The unadjusted male/female (log) wage difference increases with age. On average, young men earn 30% more than young women, middle-aged men earn 46% more than middle-aged women, and older men earn 54% more than older women. For the OLS results, the explained portion of the wage difference also increases with age. The combined effect of all the wage determinants explains 59% of the wage gap between young men and women, 65% of the wage gap between middle-aged men and women, and 68% of the wage gap between older men and women. It would appear that both the increase in the wage gap and the increase in its explained portion are due to the widening of the male/female experience and tenure gaps over time. The net effect of including housework time in the wage equation (model 2 compared to model 1) is an increase in the explained portion of 9–15 percentage points, which is within the range reported by other studies.

A more detailed comparison of the two sets of OLS results show that the gender gap in housework time, by itself, contributes 20–23 percentage points of the explained portion of the gap, 8–12 points more than that suggested by the net effect. This difference is due to housework time taking explanatory power away from (lagged) market hours and experience. The inclusion of housework time in the wage equation has a larger net effect on the explained portion of the wage gap for young workers than it does for older workers. This is because

12. An alternative explanation is that the negative effect of housework time on wages arises, in part, from not correctly controlling for current market time. If lagged market time imprecisely measures changes in market time from one year to the next, and if current housework time and current market time are negatively correlated, then a portion of housework time's negative effect on wages will be due to the bias associated with imperfectly controlling for current market hours in the wage equation. Certain groups, such as new mothers, who may be more likely to have substantial changes in their market and housework time from one year to the next, will be particularly susceptible to this type of omitted variable bias. Furthermore, because the experience variable does not include additional experience gained in the current year, the same type of bias also could be associated with the measure of experience. The authors are grateful to an anonymous referee for this point.

**TABLE 5**
Percentage of the Male/Female Wage Differential Explained by Differences in Measured Characteristics, PSID 1983–93

| Age Group | 20–34 | | | | 35–49 | | | | 50–65 | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Estimator | OLS | | HTIV | | OLS | | HTIV | | OLS | | HTIV | |
| Models | 1 | 2 | 1 | 2 | 1 | 2 | 1 | 2 | 1 | 2 | 1 | 2 |
| *Characteristics* | | | | | | | | | | | | |
| Housework (hours/week) | — | 22.6 | — | 4.2 | — | 19.7 | — | 2.2 | — | 20.9 | — | 1.9 |
| *Human capital variables (years)* | | | | | | | | | | | | |
| Education | −3.6 | −3.4 | −5.6 | −5.4 | 3.4 | 3.3 | 3.6 | 3.6 | 2.3 | 2.2 | 3.8 | 3.8 |
| Experience and experience squared | 24.9 | 20.3 | 24.2 | 23.6 | 32.2 | 26.3 | 20.6 | 20.3 | 37.7 | 28.7 | 21.1 | 20.8 |
| Tenure and tenure squared | 6.1 | 6.1 | 2.2 | 2.2 | 7.2 | 7.3 | 3.5 | 3.5 | 11.4 | 10.9 | 7.0 | 7.0 |
| Market hours (lagged) | 14.5 | 10.3 | 1.9 | 1.6 | 11.6 | 8.0 | −0.6 | −0.7 | 8.0 | 5.6 | −1.0 | −1.0 |
| Other variables | 1.8 | 2.8 | 17.3 | 17.1 | 0.1 | −0.1 | 35.5 | 35.0 | −0.3 | −0.5 | 16.9 | 16.4 |
| Total explained | 43.7 | 58.7 | 40.0 | 43.3 | 54.5 | 64.5 | 62.6 | 63.9 | 59.1 | 67.8 | 47.8 | 48.9 |
| Difference in explained portion | 15.0 | | 3.3 | | 10.0 | | 1.3 | | 8.7 | | 1.1 | |
| Male/female (log) wage differentials | 0.2902 | | | | 0.4564 | | | | 0.5366 | | | |

*Notes:* The percentages based on OLS estimates are calculated from the male/female differences in means presented in Table 1 and the estimates from OLS regressions pooled across gender but segmented by age. The percentages based on the HTIV estimates are calculated from the male/female differences in the means presented in Table 1 and HTIV regressions pooled across gender but segmented by age. Estimates from all the pooled regressions are available from the authors.

current housework time does not appear to be as correlated with previous experience for young workers as it is for older workers. These results suggest that time spent on housework today may have a feedback effect that adversely affects women's labor market experience. This may in turn amplify the adverse affect on their future wages.

Using the estimates from the HTIV procedure in the wage decomposition presents a different picture. First, although the pattern by age is similar—housework time explains more of the wage gap between young men and women than it does for older men and women—the amount explained is significantly smaller. Adding housework time only increases the explained portion of the wage gap by one to three percentage points, which is one-ninth to one-fourth the size of the effect from the OLS results. Furthermore, including housework time does not affect the portion of the gap explained by the experience variables as it does when using the OLS estimates.[13] The feedback effect, suggested by the OLS results, appears to be a function of heterogeneity bias in the returns to housework time and experience. The influence of heterogeneity bias increases with age, especially with respect to the returns to experience.

Based on the HTIV results, the authors can conclude that the male/female difference in experience is the most important determinant of the male/female wage gap; it explains 20–24% of the gap. However, the male/female housework gap does have a larger impact on the wage gap between young men and women than it does for the older age groups. For young workers, the average difference in men and women's housework time is approximately twice as important as tenure and (lagged) market hours in explaining the male/female wage differential. For middle-aged workers, it is 60% as important as tenure and education in explaining men's higher wages. Finally, for older workers, differences in time spent on housework by men and women are one-half as important as education and one-fourth as important as tenure in explaining the wage differential.

13. The authors also estimated HTIV models that treated the experience variables as endogenous. For older workers only, they were able to accept that all time-varying variables except housework time, experience, tenure, and union status were exogenous. It appears that experience and experience squared belong in the exogenous variable set for all groups except older workers. The estimates from these models differed only slightly from the ones presented in the article and are available from the authors.

## VII. CONCLUSIONS

This study extends Hersch and Stratton's research on housework and wages. Following

their approach, the authors use samples from the PSID, although for a more recent time period. The authors extend their work by expanding the analysis to include nonwhites, by disaggregating the sample into three age cohorts to determine if the effect of housework on wages has differential effects based on age, and by using an instrumental variables estimator that accounts for both heterogeneity and endogeneity in panel data.

Like Hersch and Stratton, these authors find that hours spent on housework adversely affect married women's wages and have an indeterminate effect on married men's wages. They also find evidence for women that the effect of housework on wages changes with age in a way that is consistent with the childbearing and child rearing stages of the life cycle. Each additional hour of housework time reduces the wages of young and middle-aged women by 0.1–0.4%. These results are robust even when the authors control for the number of children of different ages in the home. They find no consistent evidence that housework time affects the wage of older women, even though older women perform more housework, on average, than do younger women and as much housework, on average, as middle-aged women. The authors also find no consistent evidence that housework time affects the wages of men of any age.

It is possible that the authors have not correctly modeled the relationship between wages and housework time for the samples of men and older women. Limitations of this housework-time variable are that it does not detail the type of housework individuals are performing, and it assumes the same timing and the same intensity of effort for all types of housework activity. Knowing more about the timing and type of housework that individuals perform might shed some light on whether men and older women are more likely than are young and middle-aged women to perform less intensive housework or perform housework primarily on weekends. This is a topic for future inquiry.

These results for young and middle-aged women lead the authors to believe that it is housework time in conjunction with child care that is the true cause of the wage effects of housework time. If, as has been the case traditionally, young women are the primary care givers for small children, then it is possible that the intensity of effort put forth by young women may be much greater than the effort expended by older women and men. Child caretakers cannot postpone care until the weekends or more convenient times. The impact of housework time on wages may, in fact, be the result of the division of chores of a specific type with timing that cannot be postponed within the household.

When the analysis controls for the individual effect, the inclusion of housework time increases the explained portion of the male/female wage gap by one to three percentage points. By itself, housework time contributes 3–10% of the explained portion of the gap. Housework time has a larger net effect on the explained portion of the gap for younger workers. For older workers the OLS estimates indicate that current housework time is strongly correlated with previous experience, whereas the HTIV estimates suggest this correlation is due to the unobserved individual effect. That there may be stronger correlations between the individual effect and the experience and housework time of older workers adds support to the original conjecture that the importance of heterogeneity and endogeneity is likely to vary with age.

Unfortunately, the authors were not able to distinguish life-cycle effects from cohort effects. Some of the differences attributed to life-cycle changes may be due to changes over time that cannot be accounted for with the use of a set of year dummy variables. For example, younger women face different labor market opportunities and different societal expectations regarding working mothers than did older women. Attitudes toward the traditional distribution of housework within in the family are also changing, if numerous recent articles in the popular press are any indication. If these improvements in women's market opportunities and changes in societal attitudes cause younger couples to share housework more equally than did older couples at the same age, potential gender differences in wage/housework effects should be mitigated. Not controlling for the cohort effects might understate the effect of housework time on wages for young women and overstate it for young men relative to what was experienced by older workers at the same age. Future research will address this issue.

## APPENDIX TABLE A-1

Estimates of Housework Time and Number of Children Variables from OLS, Fixed Effects, and HTIV Models for Samples of Young and Middle-Aged Women

| | OLS | | | Fixed Effects | | | HTIV | | |
|---|---|---|---|---|---|---|---|---|---|
| **Models** | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** | **(6)** | **(7)** | **(8)** | **(9)** |
| *Young women (20–34 years old)* | | | | | | | | | |
| Housework time | −0.411*** | — | −0.387*** | −0.140*** | — | −0.129*** | −0.158*** | — | −0.146*** |
| Number of children | | | | | | | | | |
| Younger than 3 years old | — | −0.718 | 0.676 | — | −0.788 | −0.291 | — | −1.281** | −0.706 |
| 3 to 5 years old | — | −3.587*** | −2.490*** | — | −2.162** | −1.835** | — | −3.105*** | −2.693*** |
| 6 to 13 years old | — | −2.701*** | −1.968*** | — | 0.022 | 0.092 | — | −1.595** | −1.380** |
| 14 to 17 years old | — | 1.483 | 1.776 | — | 2.452 | 2.397 | — | 1.498 | 1.544 |
| *Middle-aged women (35−49 years old)* | | | | | | | | | |
| Housework time | −0.373*** | — | −0.343*** | −0.072** | — | −0.070** | −0.093*** | — | −0.091*** |
| Number of children | | | | | | | | | |
| Younger than 3 years old | — | 1.722 | 2.262 | — | 0.278 | 0.402 | — | −0.006 | 0.169 |
| 3−5 years old | — | −2.636* | −1.596 | — | −1.732 | −1.572 | — | −2.730*** | −2.496** |
| 6−13 years old | — | −2.900*** | −2.299*** | — | 1.038 | 1.116 | — | −0.440 | −0.317 |
| 14−17 years old | — | −2.040*** | −1.689** | — | 0.905 | 0.922 | — | −0.117 | −0.077 |

*Notes:* The estimates are multiplied by 100. The estimates of housework time in columns 1, 4, and 7 are from Tables 2, 3, and 4. The standard errors, the estimates of other variables, and estimates for the male and older women samples are available from the authors. Housework time is assumed to be endogenous and the fertility variables are assumed to be exogenous in the HTIV procedure. ***, **, * Significant at the 0.01, 0.05, and 0.10 levels, respectively.

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