THREE ESSAYS IN LABOR ECONOMICS: FERTILITY EXPECTATIONS AND CAREER CHOICE, SPECIALIZATION AND THE MARRIAGE PREMIUM, AND ESTIMATING RISK AVERSION USING LABOR SUPPLY DATA

A Dissertation

by

MEGAN DE LINDE LEONARD

Submitted to the Office of Graduate Studies of
Texas A&M University
in partial fulfillment of the requirements for the degree of

DOCTOR OF PHILOSOPHY

May 2007

Major Subject: Economics

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ABSTRACT

Three Essays in Labor Economics: Fertility Expectations and Career Choice, Specialization and the Marriage Premium, and Estimating Risk Aversion Using Labor Supply Data. (May 2007)

Megan de Linde Leonard, B.A., Hendrix College

Chair of Advisory Committee: Dr. Manuelita Ureta

Women, on average, are found in systematically different careers than men. The reason for this phenomenon is not fully understood, in part because expectations play a vital role in the process of career choice. Different religious groups have different beliefs on the importance of child bearing, so fertility expectations should differ by religious group. I include a woman's religious denomination in regressions on measures of occupational flexibility. Jehovah's Witnesses choose the most flexible careers followed by Pentecostal, Catholic, Baptist, and Mainline Protestant women. Jewish women generally choose the least flexible careers. This is consistent with the human capital notion that women are choosing different careers than men rather than being forced into different job paths.

If women are choosing jobs that allow them to take responsibility for home production, how does this affect their husbands? Male wage regressions that include marital status dummy variables find a marriage wage premium of 10 to 40%. This premium may occur because wives are taking responsibility for home production and husbands are free to focus their attention on productivity at work. It may also be that factors unobserved to the researcher may make a man more productive and more likely to marry. I use religious denomination as a proxy for specialization within the home. Men in more traditional religious denominations enjoy a higher marriage wage premium, which is evidence that household specialization of labor is an important cause of the wage premium.

The choice of a career, whether to marry, and most other important life decisions are dependent on one's risk tolerance. The role of risk preferences in such choices is not fully understood, largely because relative risk aversion (γ) is hard to empirically quantify. Chetty (2006) derives a formula for γ based on the link between utility and labor supply decisions. I estimate γ at the micro level using the 1996 Panel Study of Income Dynamics. I compare γ to an estimate based on hypothetical gambles and find the measures substantially different. This supports Chetty's claim that expected utility theory cannot sufficiently explain choices under uncertainty in different domains.

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CHAPTER I

INTRODUCTION

Interesting economic phenomena abound in everyday life. By the age of 5, children already assume that teachers are women and police officers are men. It isn't until later in life that one may begin to question why women are heavily concentrated in some fields and conspicuously absent from others. It may be that women are choosing careers that are amenable to the responsibilities of taking care of a family. It is also possible that they face discrimination along the career path.

One of the issues that makes this problem difficult is the importance of expectations in the career choice process. If a woman expects to have a large family and many responsibilities within the home, she may be more likely to choose a career that is compatible with this choice. Expectations are not always fulfilled and ignoring their importance can lead one to incorrect conclusions about the cause of a woman's career path. Different religious groups have different beliefs on gender roles and the importance of child bearing. Because of this, fertility expectations should differ by religious group. In Chapter II, I include a woman's religious denomination as a control in regressions on various measures of occupational flexibility. If religious denomination has significant explanatory power on women's career choices, then I can conclude that to some extent, women are choosing careers that are most compatible with their responsibilities in the home instead of being forced into a career because of limited opportunities.

If women are choosing jobs that allow them to take responsibility for home production and child rearing, how does this affect their husbands? Male wage regressions

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The journal model is *American Economic Review*.

that include marital status dummy variables almost always find that married men earn a hefty wage premium. This could be precisely because their wives are taking the primary responsibility for home production and husbands are free to focus their attention on productivity at work. It may also be that a factor unobserved to the researcher may both make a man more productive and more likely to marry. I investigate this question in Chapter III by again using religious denomination, this time as a proxy for specialization within the home. If men in more traditional religious denominations enjoy a higher marriage wage premium, this is evidence that household specialization of labor is an important cause of the wage premium.

The choice of a career, a level of specialization within the home, whether to marry and when, and most other important life decisions are heavily dependent on one's tolerance for risk. If a woman is very risk averse, for example, she may not be willing to completely specialize in home production because of the chance that her husband could die or leave. The role of risk preferences in these life choices is not fully understood, largely because risk aversion is very hard to empirically quantify. Raj Chetty (2006) uses the fact that relative risk aversion is directly related to the ratio of the income and compensated wage elasticities of labor supply and presents a formula that allows one to estimate the coefficient of relative risk aversion using labor supply data. While the primary purpose of his paper is placing an upper bound on risk aversion, Chetty's measure has the potential to be very useful in explaining individual risk preferences since most microeconomic datasets contain the necessary information to estimate risk aversion in this manner. In Chapter IV, I estimate Chetty's coefficient of relative risk aversion at the micro level using the 1996 Panel Study of Income Dynamics (PSID) and investigate his claims that the expected utility model does not do a good job of explaining choices under uncertainty in different domains.

CHAPTER II

DO FERTILITY EXPECTATIONS AFFECT WOMEN'S CAREER CHOICES?

A. Introduction

While it is likely that uttering the phrase "that's women's work" would result in nasty stares or worse, it does seem to be true that women choose systematically different careers than do men. Women are concentrated in teaching and nursing, and in academia, the humanities. Female dominated careers also tend to pay lower wages. There are different explanations for this phenomenon. It may be that women value job attributes differently than men do and are willing to pay for these attributes through lower wages. The difference between the careers of men and women could also be the result of discrimination. Women may be denied entry or advancement in some careers. The explanation for this empirical regularity is very important. If women are forced into certain careers because their market opportunities are limited, their well-being could be increased by affirmative action type policies. If women are simply responding optimally to their household's utility maximization problem, no such action is needed.

The underlying cause of the differences in careers by gender is difficult to determine, in part because expectations play such a vital role in the decision process. It may not be sufficient to look at a woman's marital status and number of children as the important factors in her career outcome. It is likely that a woman makes the decision to go to college before her decision to marry or have children. At eighteen, however, she may already know that she wants to have four children and thus decide not to be a doctor. This same woman may end up single, childless, and in a dead-end career because of her unfulfilled expectations. If one does not include her expectations in the analysis, one will come to an incorrect conclusion about the origins of her career path.

Unfortunately, most data sets employed by economists provide no information about expectations. Those that do are forced to focus on young women because the expectations that shape women's futures are formed early in life. It is then necessary to follow these women for a long period of time to see how their expectations affected their careers. Using older women could alleviate this tracking problem, but they would have to be asked about their past expectations which may be difficult to recall or separate from what actually occurred. These issues sharply narrow the data available for use in studying this problem. Sandell and Shapiro (1980) use the 1968 National Longitudinal Study of Young Women and find that expectations of higher labor force participation result in women receiving more on-the-job training. Blakemore and Low (1983) utilize data from the National Longitudinal Survey of High School Seniors in 1972 to see how expected fertility affects the choice of a college major. As things have changed a lot for women since these two surveys, this paper makes an important contribution to the literature because it brings more recent data to bear on the issue.

A proper analysis requires recent data that contains information on a woman's expectations of absence from the labor force and obligations within the home when she made her career decision. I argue that a woman's religious preference is helpful in this respect. Some religious denominations put a large emphasis on families, child rearing, and traditional gender roles. Jehovah's Witnesses, for example, believe that "a husband is the head of his family"; and "a wife should be a good helper for her husband."[1] The statement of faith of the Southern Baptist church echoes a similar sentiment; "a wife is to submit herself graciously to the servant leadership of her

[1]From http://www.watchtower.org "Family Life That Pleases God"

husband" and she "has the God-given responsibility to respect her husband and to serve as his helper in managing the household and nurturing the next generation[2]." The United Methodist Church (2004), on the other hand, has a very different view of the role of women.

> We affirm women and men to be equal in every aspect of their common life. We therefore urge that every effort be made to eliminate sex-role stereotypes in activity and portrayal of family life and in all aspects of voluntary and compensatory participation in the Church and society.

In their study of religious denomination and gender attitudes, Brinkerhoff and MacKie (1984) find that Pentecostals are among the least egalitarian denominations, followed by Baptists and Catholics. Presbyterians were the most egalitarian denomination that they investigated. Their study did not address Jews. Jews have lower fertility than other groups (Goldstein 1992), so Jewish women may expect less time out of the labor force. In contrast, devout Catholics eschew all but natural means of birth control and thus may have higher expectations of fertility than women of other religious groups.

Women know what religion they are in from a young age and thus have the opportunity to make career related decisions based on their expected fertility and family role. If the differences in the careers of men and women are due to choice in a labor market that does not discriminate against women, women in religious denominations that emphasize traditional gender roles and discourage the use of birth control should have higher fertility on average and be found in careers that are more compatible with greater in-home responsibilities. Groups that stress equality should behave in a manner more similar to women with no religious preference. If Jehovah's

[2] http://www.sbc.net/bfm/bfm2000.asp

Witness, Pentecostal, Catholic, and Baptist women choose systematically different careers than Jews and Mainline Protestant women such as Methodists and Presbyterians, then I can conclude that to some extent, women are choosing careers that are most compatible with their expected familial responsibilities instead of being forced into a career because of limited opportunities.

B. Literature Review

Economists have a lot to say about the discrimination versus choice debate. In his seminal paper, Becker (1985) posits that there are increasing returns to specialized human capital, so married couples have an incentive to create "a division of labor in the allocation of time and investment in human capital." Since women do a disproportionate amount of the family's child care and housework, they have less of their effort endowment to spend on the job and thus make less money than men with equal human capital. It is not a matter of discrimination; it is household utility maximization. McDowell's (1982) findings support the human capital approach in the careers of women in academia. Women are more likely to interrupt their careers than men, these interruptions cause some degree of knowledge obsolescence, and this obsolescence varies by field. The fact that women are more often in the humanities than the sciences is a direct response to the higher cost of career interruptions in fields where knowledge is less durable. Women aren't choosing the humanities because they cannot get jobs as scientists; they are choosing humanities because the relative costs of interruptions are lower.

Polachek's (1981) paper is very similar in spirit to McDowell's. He recognizes that variation in kinds of human capital may be as important as variations in amount. Since women do not all participate continuously in the labor force, their maximization problem should include choosing not only lifetime investment in human capital, but occupation as well. Because different occupations differ in the cost of labor force intermittency, an individual will choose the occupation that imposes the smallest penalty given a desired level of lifetime labor force participation, all else equal. He finds that there is a strong relationship between lifetime labor force participation and occupational choice. He looks at the effect of actual labor force separations on occupational choice, though theory says that expected labor force participation is the variable of interest.

It is important to note that it is a woman's expectation of labor force attachment and career path that shapes her career decisions in both the human capital and discrimination explanations. If a woman expects discrimination, she may not bother to get higher education or seek a prestigious job. Likewise, if she expects to spend a lot of time out of the labor force, she may seek a different type of job than if she planned to work continuously. Realizing the importance of expectations and the scarcity of data about it, Blau and Ferber (1991) survey college business school seniors and collect information on their career plans and earnings expectations. The women and men they survey expect the same starting salaries, but women anticipate lower earnings in later years. This holds even for women who plan on continuous employment. A woman seems to expect to make less money for performing the same job as a comparable man regardless of her planned labor force participation. This expected glass ceiling may deter her from investing in her job in the same way a man would.

Gronau (1988) also finds evidence contrary to the human capital approach. He uses simultaneous equations to trace the interrelationship between interruptions in women's careers and their wages. He finds that the skill intensity of a woman's job, which is important for earning power, is independent of her labor force plans. This seems to imply that women face restrictions in choosing their jobs.

The literature addressing this debate has produced mixed evidence. This could be in part because it is difficult to get current and general data on labor force participation plans of women, which is a key variable. In this paper I include a woman's religion as an indicator of her labor force participation plans to shed additional light on this interesting puzzle.

C. Data

I use data from the 1985 to 2001 waves of the Panel Study of Income Dynamics (PSID). The PSID is a longitudinal data set collected by the University of Michigan. The PSID includes questions on a respondent's religious preference and allows me to identify many different religious denominations. The denominations I consider are Jehovah's Witness, Pentecostal, Catholic, Baptist, Jewish, and Mainline Protestant, which includes Presbyterian, Methodist, Episcopalian, Lutheran, and "Other Protestant." Unfortunately, I only have information on the woman's current religious preference. As most individuals do not radically change their religious affiliations over time, this analysis is still expected to provide valuable information. The Encyclopedia Britannica gives statistics for religious preference over time. Out of 254,076,000 individuals in the United States in 1990, only 129,000 had converted to another religion by 1995[3].

Table 1 gives summary statistics for the women in my sample. Although the PSID is a longitudinal data set, I include each woman in the sample only once. If she appears in more that one year in the time period, her most recent information is

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[3]http://www.britannica.com/eb/article?tocId=9346285

used in the analysis[4]. I restrict the sample to include only women between the ages of 25 and 55. Women younger than 25 are not included because of the possibility that these women are combining school and work. I also exclude women who are nearing retirement. The average age in the sample is 40. Fifty-eight percent of the women are married and the average number of children under 17 years of age in the household is 1.17. Women in the sample have 12.6 years of education on average. Approximately 33% of women in the sample are Baptist, 27% are Mainline Protestant, 27% are Catholic, 3% are Pentecostal, 2% Jewish, 1% are Jehovah's Witness, and 7% have no religious preference.

D. Theory

The model used in this analysis is based on the idea of compensating differentials. Every job in the market contains many dimensions, and workers derive utility from both the wage and the amenities of a particular job. Workers are assumed to have different demands for job amenities, and firms are faced with different costs of providing the amenities. A firm will only incur costs in order to provide the desirable feature if it can then reduce the wage that it offers.

In the compensating differentials framework, if two groups of people have different preferences for wages and job amenities, utility maximization will result in different wages between the two groups and they will be found in different types of occupations (Filer 1985). The differences in the wages and occupations of men and women may simply be due to different preferences.

I must now consider why a woman might value different job attributes than would a man and what amenities she might demand. Women differ from men in

[4]Results are not sensitive to the year in which a woman appears in the data set.

Table 1. Summary Statistics from PSID-Women Only

Variable	Obs	Mean	Std. Dev.	Min	Max
Age	9508	40.191	9.289	25	55
Number of Children	9508	1.171	1.251	0	9
Married	9508	0.582	0.493	-	-
Years of Education	9187	12.590	2.620	1	17
Northeast	7105	0.153	0.360	-	-
Northcentral	7105	0.224	0.417	-	-
South	7105	0.455	0.498	-	-
West	7105	0.166	0.372	-	-
Proportion Jehovah's Witness	9508	0.010	0.101	-	-
Proportion Pentecostal	9508	0.033	0.178	-	-
Proportion Catholic	9508	0.265	0.441	-	-
Proportion Baptist	9508	0.329	0.470	-	-
Proportion Mainline Protestant	9508	0.271	0.445	-	-
Proportion Jewish	9508	0.018	0.133	-	-
Proportion with No Religious Preference	9508	0.073	0.261	-	-

Notes: Data is from the PSID and includes only women between the ages of 25 and 55.

that they generally have the primary responsibility for child care. Children require constant care, so if a woman chooses to work outside the home, the family must find and finance child care and transport the child there every day. Because leaving children in another's care is costly both in monetary and psychic terms, there is value in having a job that is compatible with the responsibilities of having children. Sick children cannot go to school or daycare, so someone must stay home with children when they are ill as well. Having a job that allows a mother to deal with unforseen circumstances is desirable.

The job amenity that I consider is flexibility. According to a survey done in 1997 by the Pew Research Center[5], 73% of mothers said that job flexibility was a very important job attribute. I assume that women value jobs that allow them to take time off if they need it, arrange their schedules around the schedules of their children, etc. The more children a woman has, the more time she will need off for maternity leave, the more likely it is that there will be an illness on a given day, and the more schedules there are to juggle. I therefore assume that the more children a woman expects to have, the more she will value a flexible job.

As shown in Figure 1, women who expect many children have steeper indifference curves indicating that they are more willing to trade off wages for flexibility. They work at firms that have a lower cost of providing flexibility than men or women who expect to have few children. Women with very steep indifference curves may locate themselves on the boundary, consuming the maximum amount of flexibility and accepting a wage of zero, i.e., not working in the market.

[5]http://people-press.org/reports/display.php3?PageID=520

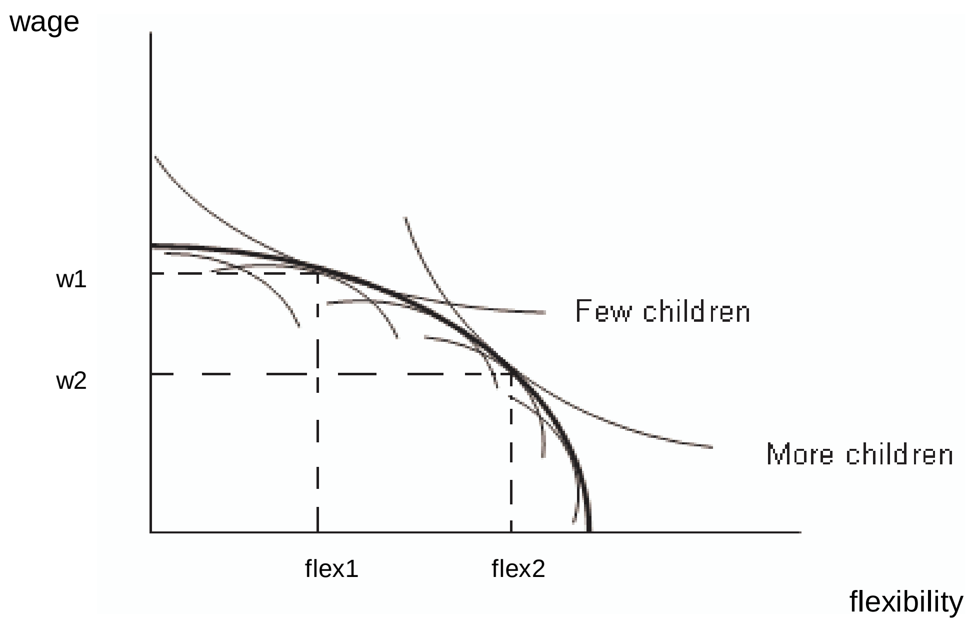

Fig. 1. Wage-flexibility Trade-off: Compensating Differentials Framework

The wage-flexibility trade-off probably explains why we see so many women accepting relatively low wages to become teachers. Teaching is attractive to women because it allows a mother to take advantage of her school age children's predictable schedules. Teaching does not, however, offer a great deal of flexibility in regards to temporary shocks such as a child's illness. Because flexibility has more than one dimension that may be important to mothers, I define two components of flexibility: predictable flexibility and unpredictable flexibility.

Jobs have a great deal of predictable flexibility if they are very compatible with the day-to-day family needs that can be anticipated. Unpredictable flexibility refers to the ability to accommodate unexpected events such as a child's illness or a forgotten school project. When children are young, the majority of their needs may be unpredictable, making it desirable for the mother to stay at home for a period of time. Jobs that easily accommodate entry and exit from the labor force would be considered unpredictably flexible, for example.

E. Findings

Flexibility is difficult to define empirically. Because of this, I will explore several different empirical definitions of flexibility and see if the estimated effect of religious preference on measured job flexibility is consistent across the different flexibility measures. As described above, flexibility has more than one dimension. Mothers value jobs that allow them to be home for their children's predictable needs and jobs that allow them to be home in the case of unexpected shocks. I construct one measure of predictable flexibility, two measures of unpredictable flexibility, and one measure of general flexibility.

The predictable components of job flexibility are those that allow workers to take advantage of their child's schedule. I construct one variable that represents predictable flexibility empirically: the average hours worked in a usual week in the occupation. Long hours at work allow for less time with the family and require more money to be spent on child care. This variable is calculated from the 1990 Census 1% sample for each of the 505 occupation codes using data on men only. Women are not included in calculating job attributes to avoid endogeneity problems.

The two measures of unpredictable flexibility are the fraction of men who work from home in each of the 505 occupations and the rate of depreciation for each of 17 occupational classes. In order to work in the market and also provide for her children's needs, a woman may choose to work from home if the option is available to her. The 1990 Census asks respondents how they get to work each day. A little over 36,000 of the respondents report that they work from home. Since the ability to work from home allows women to both work and care for their children in many cases, this measure is an important piece of the flexibility puzzle. I use the fraction of men who work from home in a given occupation as a measure of each job's flexibility.

Instead of working from home while her children are young, a woman may decide to drop out of the labor force for a period of time. It would then be necessary for her to consider the amount of depreciation of job related skills associated with a given career. The larger is the job related knowledge depreciation, the more costly is any interruption in participation. If a woman expects interruptions to be long, she may opt for a career with less skills depreciation.

I estimate rates of human capital depreciation for 17 occupational categories using Chutubtim's (2005) modification of Ureta and Welch (2001). Their spin on the standard Mincer wage regression is to differentiate workers with the same amount of work experience who accumulated this experience at different points in their careers. Since women have more career interruptions than men, acknowledging that older work experience may have different returns than more recent experience is particularly important. For women who work continuously, this specification is the standard Mincer model.

Following Ureta and Welch (2001), if a person works full time between times t_1 and t_2, the instantaneous rate of human capital accumulation at time $t = \tau$ is

$$\frac{dlnK(\tau)}{d\tau} = \alpha \qquad\qquad \text{for } t_1 \leq \tau \leq t_2.$$

Let δ denote the instantaneous rate at which accumulated human capital depreciates from the time at which is it acquired, τ, to the present, T:

$$\text{Depreciation by time } T \text{ of human capital attained at } \tau = \int_{\tau}^{T} \alpha\delta \, dt = \alpha\delta(T - \tau).$$

Net human capital accumulation at T for work done between t_1 and t_2 is:

$$\begin{aligned} \int_{t_1}^{t_2} (\alpha - \alpha\delta(T - t)) \, dt &= \alpha(t_2 - t_1) - \alpha\delta \left(T(t_2 - t_1) - \frac{t_2^2 - t_1^2}{2} \right) \\ &= \alpha D - \alpha\delta D \left(T - \frac{t_2 + t_1}{2} \right) \end{aligned}$$

$$= \alpha D - \alpha\delta(DA),$$

where $D = t_2 - t_1$ is the duration of the employment spell and $A = T - \frac{t_2+t_1}{2}$ is the average age of the employment spell.

For a worker with I spells, net accumulation of human capital is:

$$\alpha\left(\sum_{i=1}^{I} D_i\right) - \alpha\delta\left(\sum_{i=1}^{I} D_i\overline{A}\right)$$

where $\sum_{i=1}^{I} D_i$ is the aggregate spell duration and $\overline{A}$ is the average age of the I spells.

In order to estimate the rate of depreciation for various occupations, I use the following wage equation:

$$\ln W_i = \beta_0 + \beta_1 X_1 + \alpha\sum_{t=1}^{T} D_{i,t} + \alpha\delta\left(\sum_{t=1}^{T} D_{i,t}\overline{A_i}\right) + \varepsilon_i \tag{2.1}$$

where $\ln W_i$ is the log of the weekly wage and X includes indicator variables for living in a rural area, living in the center of a metropolitan area, living in the suburbs, married, education level, race, region, whether the youngest child is of school age, whether the woman speaks English, and whether she works full-time. Also included in the X vector is the number of children in the woman's household and the number of years since her highest degree was earned.

Census data does not contain information on a worker's complete employment history. Because of this, I use the method in Chutubtim (2005) to construct values for $\sum_{i=1}^{I} D_i$ and $\overline{A}$. By assuming that women only leave the labor market in order to have children, I can construct employment histories for the women in the sample using information on their age, education, number of children, and age of children.

Chutubtim makes the reasonable assumption that time out of the labor force will differ by education level and estimates the maximum career interruption for each education category. For each level of education she is interested in three probabilities, the percentage of women who do not take time out of the labor force after giving birth ($P1$), the percentage of women who take 3 years out of the labor force ($P2$), and the percentage of women who take 12 years out of the labor force ($P3$). To simplify the analysis, she considers women who have only one child. To calculate these probabilities, one needs to know how much time a woman is out of the labor force after the birth of a child. The Census data does not directly contain this information, but a woman's occupational information allows one to draw some inferences. If a woman is out of the labor force in 1990 but has a reported occupation, she has been out of the labor force for no more than 5 years. If a woman is out of the labor force in 1990 and she does not have a reported occupation, one can infer that she has been out of the labor force for more than 5 years.

Using this information, $P1$ is assumed to be the proportion of women with a child under the age of one who are in the labor force. $P2$ is the proportion of women with a child between the ages of 1 and 5 who are out of the labor force for fewer than five years, and $P3$ is the proportion of women with a child between the ages of 6 and 17 who are out of the labor force for more than 5 years. For each level of education, the maximum number of years of career interruption is given by $(P1 * 0) + (P2 * 3) + (P3 * 12)$. Chutubtim calculates these probabilities from the 2000 Census and finds that for women with less than a high school education, the maximum time out of the labor force per child is 9 years. High school graduates have a maximum time out of 7 years, women with some college 4 years, and college graduates 2 years.

Chutubtim assumes that during the n^{th} year of a career interruption spell, the probability that a woman will return to work equals the labor force participation rate of women with one child age n. These probabilities can be found in Table 2 and are taken directly from Chutubtim (2005). On the birth of her first child, a high school drop out will return to the labor force after zero years out with probability 35.06%. With probability 24.89%, she will be assigned 1 year of career interruption. With probability 15.99% she will be assigned 2 years out of the labor force, with probability 9.63% she will be assigned 3 years, and so on until 100% of women return after the 9^{th} year.

I use the probabilities in Table 2 to assign career interruption spells to all women aged 25 to 44 in the 1990 Census based on their educational attainment, number of children, and children's ages. For example, 71.82% of women who have at least a college education are randomly assigned zero years out of the labor force upon the birth of their first child. 20.09% of women with at least a college education are assigned an interruption of 1 year when their first child is born. The remaining 8.09% of college educated women are assigned 2 years out of the labor force at the birth of their first child. If a woman has more than one child, she may have several career interruption spells. Take, for example, a woman with two children. If the difference in her children's ages is greater than her first assigned career interruption spell, she will be assigned an additional spell using the probabilities in Table 2. If the initial spell is greater than the age gap between her children, she will have both children during the initial interruption and return to work permanently after the first spell.

After career interruption spells have been assigned, I can easily calculate experience for each woman based on her assigned career interruptions. Given the information in the Census, I can also define potential experience for each woman, i.e., age minus education minus 6. For women who are college graduates, the correlation between potential experience and imputed experience based on assigned interruptions is 98.73%. For high school graduates, the correlation is 97.2%. For high school dropouts, on the other hand, the correlation between potential experience and assigned experience is 81.42%. High school dropouts have significantly more children than the rest

Table 2. Percentage of Women Returning to the Labor Force Each Year after the Birth of a Child

Education level	Child's age	Percentage of women returning to the labor force
Less than high school	0	35.06%
	1	24.89%
	2	15.99%
	3	9.63%
	4	6.47%
	5	3.70%
	6	2.09%
	7	1.20%
	8	0.51%
	9	0.47%
	Total	100%
High school	0	55.89%
	1	25.59%
	2	11.55%
	3	4.52%
	4	1.63%
	5	0.57%
	6	0.18%
	7	0.07%
	Total	100%
Some college	0	67.41%
	1	22.21%
	2	7.60%
	3	2.11%
	4	0.67%
	Total	100%
College or higher	0	71.82%
	1	20.09%
	2	8.09%
	Total	100%

Notes: This information comes from Table 6 of Chutubtim (2005) and is calculated using the 2000 Census. The percentage of women that return to the labor market in the n^{th} year of a career interruption spell is assumed to be equal to the labor force participation rate of women with one child age n.

of the sample, so it is probable that potential experience is a less accurate measure of experience for these women as they have likely taken time out of the labor force since their last year of schooling.

I estimate equation 2.1 for 17 occupational categories based on the Census occupation codes: Executive, administrative, and managerial occupations, Professional speciality occupations, Teachers and school workers, Athletes and artistic occupations, Technicians, Sales occupations, Office supervisors, Administrative support occupations, Clerks, Communications operators, dispatchers, and investigators, Food service occupations, Health service occupations, Cleaning service occupations, Miscellaneous service occupations, Farming, forestry, and fishing occupations, Precision production, craft, and repair occupations, and Operators, fabricators, and laborers. The categories were as finely partitioned as possible while still including reasonable numbers in each group.

The estimated rate of depreciation for each occupational group is found by dividing the coefficient on duration-weighted average age of the employment spells by the aggregate employment duration, i.e., $\frac{\alpha\delta}{\alpha}$. The regression results from equation 2.1 can be found in Table 3. The perverse signs on experience for cleaning and food service occupations are driven by a small number of women with low wages and no children and thus very large values of imputed experience. The imputed measures are more noisy for these women.

The depreciation rates and their standard errors for each of the 17 occupational categories can be found in Table 4. The ranking of occupations by human capital depreciation seems reasonable. Food service occupations have a low level of depreciation, 3.9% per year. Administrative support occupations, clerks, production, craft, and repair occupations, communications operators, and cleaning occupations all have depreciation rates of around 4.5%. Technicians and professional speciality occupa-

Table 3. Estimation of Equation 2.1: Log Annual Wage

Variable	Coefficient	Std. Err
Aggregate employment duration		
Cleaning service occupations	-0.012***	0.002
Athletes and artistic occupations	0.022***	0.002
Teachers and school workers, except post secondary	0.027***	0.001
Office supervisors	0.031***	0.002
Communications operators,dispatchers, and investigators	0.023***	0.001
Sales occupations	0.012***	0.001
Executive, administrative, and managerial occupations	0.043***	0.001
Precision production, craft, and repair occupations	0.024***	0.002
Administrative support occupations	0.012***	0.001
Operators, fabricators, and laborers	0.011***	0.001
Health service occupations	0.003*	0.002
Professional speciality occupations	0.064***	0.001
Farming, forestry, and fishing occupations	-0.005*	0.003
Clerks	0.01***	0.001
Technicians	0.043***	0.001
Food service occupations	-0.012***	0.001
Miscellaneous service occupations	0.001	0.002
Duration-weighted average age of employment spells		
Cleaning service occupations	0.001***	0.000
Athletes and artistic occupations	-0.001***	0.000
Teachers and school workers, except post secondary	-0.001***	0.000
Office supervisors	-0.002***	0.000
Communications operators,dispatchers, and investigators	-0.001***	0.000
Sales Occupations	-0.001***	0.000
Executive, administrative, and managerial occupations	-0.002***	0.000
Precision production, craft, and repair occupations	-0.001***	0.000
Administrative support occupations	-0.001***	0.000
Operators, fabricators, and laborers	-0.001***	0.000
Health service occupations	0.000	0.000
Professional speciality occupations	-0.004***	0.000
Farming, forestry, and fishing occupations	0.000	0.000
Clerks	0.000***	0.000
Technicians	-0.002***	0.000
Food service occupations	0.000***	0.000
Miscellaneous service occupations	0.000***	0.000
Does not speak English	-0.143***	0.008
Black	-0.019***	0.003
Hispanic	-0.042***	0.004
Number of Children	-0.047***	0.001
Married	0.014***	0.002
Youngest child of school age	-0.068***	0.003
Center of a metropolitan area	0.11***	0.003
Suburbs	0.123***	0.002
Rural	-0.078***	0.002
High school graduate	0.116***	0.004
Some college	0.237***	0.004
College graduate	0.498***	0.005
Years since last degree	0.006***	0.001
Full time work	0.717***	0.003
Constant	4.651***	0.008
Observations	261,252	
F-statistic	3570.74	

Notes: *Indicates significance at the 10% level, ** 5%, *** at the 1% level. Data comes from the 1990 Census 1% sample. Controls for region are included but not shown.

Table 4. Estimated Depreciation Rates by Occupation

Occupation	Obs	Depreciation	Std. Err
Miscellaneous service occupations	8879	0.252	0.200
Farming, forestry, & fishing occupations	1637	0.005	0.045
Food service	12069	0.039	0.006
Administrative Support Occupations	36365	0.044	0.004
Clerks	17766	0.044	0.006
Precision production, craft, & repair occupations	6540	0.044	0.004
Communications operators,dispatchers, & investigators	11965	0.045	0.003
Cleaning service occupations	5152	0.046	0.006
Teachers & school workers, except post secondary	21461	0.048	0.003
Office supervisors	5053	0.048	0.004
Executive, Administrative, & Managerial Occupations	34409	0.054	0.001
Operators, fabricators, & laborers	24517	0.055	0.004
Athletes & artistic occupations	4242	0.055	0.006
Technicians	12159	0.058	0.002
Sales Occupations	25344	0.059	0.005
Professional speciality occupations	24849	0.063	0.001
Health Service Occupations	8850	0.075	0.025
Total	261257		

Notes: All depreciation rates are significant except Miscellaneous service and Farming, forestry, and fishing occupations. These values are calculated using the 1990 Census 1% sample using Chutubtim's (2005) modification of Ureta and Welch (2001). The depreciation rates are calculated by dividing the coefficient on duration-weighted average age of employment spells by the coefficient on aggregate employment duration in the wage regression given in Table 3 for each of the 17 occupational categories.

tions have depreciation rates of 5.8 and 6.3%, respectively. These careers generally require a large amount of training, job-specific knowledge, and education so large depreciation estimates are reasonable. Health service occupations have the largest point estimate for depreciation, but it is less precisely estimated than the aforementioned occupational categories. Because the depreciation rate for miscellaneous service occupations and farming, forestry, and fishing occupations is not significantly different from zero, they will be assigned a value of zero for this flexibility measure.

One way to lessen the wage penalty due to human capital depreciation is to continue to work. Many women find that working part time allows them to keep up in their field while mitigating the household costs associated with being a working mother. According to the Pew Research Center survey, 44% of mothers of children under 18 say that they would prefer to work part time, compared to only 30% who would prefer to work full time. Working part time allows a mother to arrange her schedule around her child's schedule as well as have a greater ability to respond to unexpected needs. Because of the apparent widespread appeal of working part time, this is included as a general measure of flexibility. This measure is defined by asking: given that a man is in occupation i, what is the probability that he works part time?

Table 5 gives the values of the flexibility variables for selected occupations. The results accord with expectations; child care workers and food service workers work relatively few hours on average while veterinarians and farmers work long hours. As expected, no air traffic controllers or telephone lineman work from home. Dressmakers, painters, and authors do frequently work from home, however. According to the depreciation measures, waitresses and maids have flexible jobs while physicians and chemical engineers do not. Twenty eight percent of teacher aides work part time while less than 0.5% of physicists and astronomers do.

I am proposing that religious preference gives us information on expected fertil-

Table 5.: Flexibility Measures for Selected Occupations

Most Flexible		Least Flexible	
Average Usual Hours Worked			
Child care workers, private household	22.4	Members of armed forces	50.588
Library attendants & assistants	33.4	Veterinarians	53.273
Musicians & composers	35.161	Officers, pilots, & pursers; ship	55.196
Food service workers	38.344	Farmers (owners & tenants)	57.155
Fraction who work from home			
Dressmakers & seamstresses, except factory	0.156	Postmasters & mail superintendents	0
Child care workers, private household	0.2	Marine scientists	0
Painters & sculptors	0.227	Telephone linemen & splicers	0
Authors	0.423	Air traffic controllers	0
Depreciation of Human Capital			
Waiters & waitresses	0.039	Securities & financial services sales	0.059
Typists	0.044	Chemical engineers	0.063
File clerks	0.044	Physicians	0.063
Maids & housemen	0.046	Health aides	0.075
Fraction who work part time			
Education teachers	0.167	Policemen & detectives	0.005
Musicians & composers	0.211	Physicists & astronomers	0.005
Teacher aides, except school monitors	0.279	Petroleum engineers	0.006
Child care workers, private household	0.6	Members of armed forces	0.008

Notes: Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text.

Table 6. Mean Number of Children by Religious Preference

Variable	Obs	Mean	Std. Err	Std. Dev.	Min	Max
Jehovah's Witness	98	1.520	0.158	1.568	0	7
Pentecostal	310	1.332	0.071	1.247	0	6
Catholic	2521	1.311	0.026	1.329	0	9
Baptist	3131	1.224	0.022	1.251	0	9
None	698	1.082	0.047	1.242	0	6
Mainline Protestant	2578	0.986	0.022	1.138	0	6
Jewish	172	0.808	0.083	1.083	0	5

Notes: Data is from the PSID and includes women between the ages of 25 and 55 only.

ity. While not all expectations will be realized, if the proposed relationship holds in general, religious preference should be related to number of children. Table 6 shows that this is indeed the case. Jehovah's Witnesses have more children than any other group with an average of 1.5 children. Pentecostal and Catholic women have the next largest families, 1.3 children on average. Baptist women have 1.2 children on average, significantly more than women professing no religious preference. Mainline Protestant women have an average of 0.99 children. As expected, Jewish women have the fewest children with an average of 0.81.

Women who expect to have many children have less incentive to invest in education than women who expect to have few children because time out of the labor force means less time to benefit from investments in education. Table 7 examines years of education by religious preference. The results are consistent with group ideology and with actual fertility as seen in Table 6. Jehovah's Witnesses, a group that stresses traditional gender roles and has the highest fertility on average, get significantly less

Table 7. Mean Years of Education by Religious Preference

Variable	Obs	Mean	Std. Err	Std. Dev.
Jehovah's Witness	93	11.376	0.246	2.372
Pentecostal	297	11.848	0.136	2.342
Catholic	2421	12.057	0.067	3.287
Baptist	3030	12.353	0.038	2.089
Mainline Protestant	2522	13.189	0.046	2.304
None	661	13.195	0.096	2.470
Jewish	163	15.252	0.161	2.056

Notes: Data is from the PSID and includes women between the ages of 25 and 55 only.

education than any other group. Pentecostal women, another very conservative group with high fertility, get the next fewest years of education, less than a high school degree on average. Catholic women have more children on average than Baptist women and get significantly less education. Average years of education are nearly identical for women with no religious preference and the more socially liberal Mainline Protestant denominations. Jewish women, the group with the lowest fertility, get more than 2 years more education on average than any other group.

The theory predicts that women who expect to have more children or who anticipate having the primary responsibility for housework should choose more flexible careers. To test this theory, I regress various measures of flexibility on indicator variables for a woman's religious preference. The PSID contains information on each individual woman's occupation and she is assigned each of the four measures of flexibility based on that occupation. I then estimate the following equation using ordinary least squares (OLS):

$$\begin{aligned} Flexibility_i &= \alpha + \beta_1 Jehovah'sWitness_i + \beta_2 Pentecostal_i \\ &\quad + \beta_3 Catholic_i + \beta_4 Baptist_i + \beta_5 MainlineProtestant_i \\ &\quad + \beta_6 Jewish_i + \varepsilon_i. \end{aligned} \tag{2.2}$$

where i indexes an individual woman and flexibility is defined for a given occupation by four different measures: the average usual hours worked per week by men in the occupation, the fraction of men in the occupation who work from home, depreciation, and the fraction of men in the occupation who work part time. If a woman does not work, she is assigned a value of 1 for fraction who work part time and fraction who work from home. Women who do not work cannot be assigned a value for average usual hours worked or depreciation, so there are fewer observations in these regressions. The results are not sensitive to dropping all women who do not work.

Because two of the flexibility measures lie between 0 and 1, the variance of ε will depend on the values of the right-hand-side variables. To deal with this problem, I use robust standard errors in all regressions. The predicted values of the flexibility measures for these two measures never lie outside the unit interval, so this drawback of OLS is not a concern.

Table 8 shows the results of the regressions of the various flexibility variables on indicator variables for religion. When flexibility is defined by fraction who work from home and fraction part time, positive coefficients denote more flexible careers. In the depreciation and usual hours worked specifications, negative coefficients signify more flexibility. The omitted group is women with no religious preference.

Religious ideology, actual fertility, and years of education all suggest that if

Table 8. Religion as a Predictor of Career Flexibility

	Flexibility Measure			
	Fraction part time	Fraction work from home	Average hours worked	Depreciation
Jehovah's Witness	0.138**	0.129**	-1.789***	-0.004*
	(0.04)	(0.05)	(0.45)	(0.00)
Pentecostal	0.101***	0.096***	-0.837**	-0.001
	(0.03)	(0.03)	(0.28)	(0.00)
Catholic	0.109***	0.107***	-0.189	-0.001
	(0.02)	(0.02)	(0.17)	(0.00)
Baptist	0.042*	0.038*	-0.754***	-0.001
	(0.02)	(0.02)	(0.16)	(0.00)
Mainline Protestant	0.020	0.018	-0.119	-0.000
	(0.02)	(0.02)	(0.17)	(0.00)
Jewish	0.018	0.021	0.979**	0.005***
	(0.04)	(0.04)	(0.33)	(0.00)
Constant	0.219***	0.207***	43.123***	0.049***
	(0.02)	(0.02)	(0.15)	(0.00)
Observations	9508	9508	7202	7202
F-statistic	14.319	13.683	14.306	3.882

Notes: *Indicates significance at the 10% level, ** at the 5%, *** at the 1% level. Standard errors in parenthesis. Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text. The individual level data is from the PSID for women between the ages 25 and 55.

women are choosing careers in order to balance work and family, Jehovah's Witnesses will choose the most flexible careers followed by Pentecostal, Catholic, and Baptist women. Mainline Protestant women and Jewish women are not expected to seek more flexible careers than women with no religious preference. The hypothesis is supported in general across all definitions. Jehovah's Witnesses, a group that emphasizes traditional gender roles and the group with the largest number of children on average, are found in significantly more flexible careers than women with no religious preference regardless of the definition of flexibility. The magnitude of the coefficients is also largest for these women in every specification. Pentecostal women are found in significantly more flexible careers as well. Catholic women have the next largest coefficient magnitudes in general, followed by Baptist women. This ordering is consistent with religious ideology, actual fertility, and years of education. Because the Mainline Protestant groups tend to emphasize equality, the hypothesis suggests that these women will not choose significantly more flexible careers than women with no religious preference. This is indeed the case; while the coefficients suggest a preference for more flexibility, none of them are statistically significant. Jewish women are found in careers with a higher level of depreciation and a greater number of hours worked on average than women with no religious preference. Results are not sensitive to the inclusion of age as a control variable.

F. Further Investigation of the Channels through Which Religion Affects Career Outcomes

Women in different religious groups clearly differ in their tastes for job flexibility. As demonstrated earlier, women in more conservative religions have more children and get fewer years of education on average. I argue that these differences are due to ideological differences by religion in the function of the family and the role of women. Women in religions that stress child-rearing as the primary responsibility of the wife are likely to have more children and less likely to invest in education. I would like to know how these two variables affect the choice of career flexibility and whether religious denomination has explanatory power on job flexibility even after controls for these channels have been added.

While affecting the choice of education and number of children are the most obvious ways that religious denomination affects outcomes, denomination may have an effect on career choice even controlling for these factors. Not all expectations are fulfilled, and religious denomination is a proxy for unfulfilled expectations even when controls for children and education are added. Denomination may also pick up differences in responsibilities in the home.

In Table 9 I examine the impact of religious denomination on career flexibility when I control for number of children, the most obvious channel by which religion effects career choice. Specifically, I estimate the following equation:

$$\begin{aligned} Flexibility_i &= \alpha + \beta_1 Jehovah'sWitness_i + \beta_2 Pentecostal_i \\ &\quad + \beta_3 Catholic_i + \beta_4 Baptist_i + \beta_5 MainlineProtestant_i \\ &\quad + \beta_6 Jewish_i + \beta_7 Kids_i + \varepsilon_i \end{aligned} \tag{2.3}$$

where i indexes individual women in the PSID between the ages of 25 and 55 and $Kids_i$ is the number of children under 17 in the household. While number of children significantly increases the level of job flexibility, the magnitude of the coefficient is smaller than the coefficient on religious denomination for most groups in every specification. Comparing these results to those in Table 8, we see that controlling for

Table 9. Religion as a Predictor of Career Flexibility with Control for Number of Children

	Flexibility Measure			
	Fraction part time	Fraction work from home	Average usual hours worked	Depreciation
Jehovah's Witness	0.122**	0.113*	-1.709***	-0.004
	(0.04)	(0.04)	(0.45)	(0.00)
Pentecostal	0.092**	0.087**	-0.802**	-0.001
	(0.03)	(0.03)	(0.28)	(0.00)
Catholic	0.101***	0.099***	-0.166	-0.001
	(0.02)	(0.02)	(0.17)	(0.00)
Baptist	0.037*	0.033	-0.721***	-0.001
	(0.02)	(0.02)	(0.16)	(0.00)
Mainline Protestant	0.024	0.022	-0.128	-0.000
	(0.02)	(0.02)	(0.17)	(0.00)
Jewish	0.028	0.031	0.930**	0.005**
	(0.03)	(0.04)	(0.33)	(0.00)
Number of children	0.036***	0.037***	-0.163***	-0.001**
	(0.00)	(0.00)	(0.04)	(0.00)
Constant	0.180***	0.168***	43.285***	0.049***
	(0.02)	(0.02)	(0.15)	(0.00)
Observations	9508	9508	7202	7203
F-statistic	29.058	28.273	15.336	4.831

Notes: * Indicates significance at the 10% level, ** at the 5%, *** at the 1% level. Standard errors in parenthesis. Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text. The individual level data is from the PSID for women between the ages 25 and 55.

actual fertility has very little effect on the coefficients on religious denomination.

Controlling for education has a greater effect on the magnitude of the coefficients on religious denomination than does controlling for number of children, as shown in Table 10. Education is strongly negatively related to job flexibility. This is to be expected since wages and flexibility are inversely related and education and wages are positively related. Flexibility decreases as education increases. High school dropouts are found in the most flexible careers. College graduates in less flexible careers than high school graduates and those with some college.

It appears that the primary way in which religious denomination affects behav-

Table 10. Religion as a Predictor of Career Flexibility with Control for Education

	Flexibility Measure			
	Fraction part time	Fraction work from home	Average usual hours worked	Depreciation
Jehovah's Witness	0.074	0.064	-1.359*	-0.002
	(0.05)	(0.05)	(0.53)	(0.00)
Pentecostal	0.054	0.049	-0.515	0.000
	(0.03)	(0.03)	(0.29)	(0.00)
Catholic	0.080***	0.078***	-0.063	-0.000
	(0.02)	(0.02)	(0.17)	(0.00)
Baptist	0.007	0.003	-0.529***	0.000
	(0.02)	(0.02)	(0.15)	(0.00)
Mainline Protestant	0.022	0.020	-0.086	-0.000
	(0.02)	(0.02)	(0.16)	(0.00)
Jewish	0.080*	0.083*	0.574	0.003**
	(0.03)	(0.03)	(0.35)	(0.00)
High school graduate	-0.182***	-0.184***	0.463***	0.002***
	(0.01)	(0.01)	(0.12)	(0.00)
Some college	-0.251***	-0.254***	0.920***	0.004***
	(0.01)	(0.01)	(0.13)	(0.00)
College graduate	-0.306***	-0.306***	1.712***	0.008***
	(0.01)	(0.01)	(0.16)	(0.00)
Advanced degree	-0.272***	-0.276***	1.506***	0.007***
	(0.02)	(0.02)	(0.16)	(0.00)
Constant	0.418***	0.409***	42.228***	0.044***
	(0.02)	(0.02)	(0.17)	(0.00)
Observations	9508	9508	7202	7202
F-statistic	66.669	64.678	24.387	24.074

Notes: * Indicates significance at the 10% level, ** at the 5%, *** at the 1% level. Standard errors in parenthesis. Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text. The individual level data is from the PSID for women between the ages 25 and 55.

ior and thus career choice is through investments in education. It is possible that certain religious groups stress the importance of education more than others for reasons unrelated to fertility, but even when controls for education are added, religious denomination still has significant explanatory power over career choice. Jehovah's Witness and Baptist women are less likely to work in careers that require long hours. Catholic women are significantly more likely to work part time or work from home than women with no religious preference. When controls for education are added, Jewish women are more likely to work part time or work from home than women with no religious preference.

While it is easy to argue that religious preference has a causal effect on number of children and years of education, it is possible that there are characteristics that are correlated with a woman's religion and the flexibility of her job for reasons unrelated to fertility expectations. It is likely that different religious denominations are more prevalent in some regions than others. Since flexibility could potentially have regional characteristics as well, I add controls for region to equation 2.2. The PSID provides information on which of the four main regions the respondent lives: Northeast, North central, West, or South. South is the omitted category. Marital status is included as a control because women who are not married may have more (or less) freedom in choosing their occupation. Table 11 reports the estimates of the flexibility regressions that include education, number of children, region, and a dummy variable indicating whether the woman is married in the flexibility regression. Married women are more likely to work part time and work from home than are unmarried women. Region does not appear to have a strong impact on job flexibility, but women in the Northeast are in jobs with higher rates of depreciation than are women in the South.

Jehovah's Witness women are significantly more likely to be in jobs that require fewer usual hours worked. With the inclusion of control variables, the significance of

Table 11. Religion as a Predictor of Career Flexibility with Control Variables Added

	Flexibility Measure			
	Fraction part time	Fraction work from home	Average usual hours worked	Depreciation
Jehovah's Witness	0.055	0.044	-1.021*	-0.002
	(0.05)	(0.05)	(0.44)	(0.00)
Pentecostal	0.047	0.042	-0.466	0.001
	(0.03)	(0.03)	(0.28)	(0.00)
Catholic	0.078***	0.077***	-0.121	-0.001
	(0.02)	(0.02)	(0.17)	(0.00)
Baptist	0.013	0.008	-0.386*	0.001
	(0.02)	(0.02)	(0.16)	(0.00)
Protestant	0.012	0.010	-0.020	-0.000
	(0.02)	(0.02)	(0.17)	(0.00)
Jewish	0.052	0.055	0.802*	0.002
	(0.03)	(0.03)	(0.40)	(0.00)
Kids	0.036***	0.037***	-0.165***	-0.000*
	(0.00)	(0.00)	(0.03)	(0.00)
Married	0.036***	0.036***	0.142	0.000
	(0.01)	(0.01)	(0.08)	(0.00)
Northeast	0.025	0.022	-0.021	0.002*
	(0.01)	(0.01)	(0.14)	(0.00)
North central	-0.013	-0.016	0.147	0.001
	(0.01)	(0.01)	(0.11)	(0.00)
West	0.023	0.024	0.212	-0.000
	(0.01)	(0.01)	(0.12)	(0.00)
High school graduate	-0.179***	-0.182***	0.411***	0.003***
	(0.01)	(0.01)	(0.12)	(0.00)
Some college	-0.242***	-0.246***	0.844***	0.005***
	(0.01)	(0.01)	(0.13)	(0.00)
College graduate	-0.294***	-0.294***	1.734***	0.009***
	(0.01)	(0.01)	(0.16)	(0.00)
Advanced degree	-0.240***	-0.244***	1.289***	0.008***
	(0.02)	(0.02)	(0.18)	(0.00)
Constant	0.350***	0.342***	42.215***	0.043***
	(0.02)	(0.02)	(0.19)	(0.00)
Observations	9359	9359	6925	6925
F-statistic	61.218	59.999	16.867	18.454

Notes: * Indicates significance at the 10% level, ** at the 5%, *** at the 1% level. Standard errors in parenthesis. Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text. The individual level data is from the PSID for women between the ages 25 and 55.

Table 12. Means by Religion for Age at First Birth.

Variable	Obs	Median	Mean	Std. Dev	Min	Max
Baptist	1046	20	21.478	4.368109	13	38
Lutheran	412	22	22.617	4.52498	13	41
Mainline Protestant	836	22	23.428	4.733711	12	42
Catholic	1617	22	23.559	5.112025	13	42
Jewish	127	25	26.15	5.072614	17	39

Notes: Data is from the 1998 General Social Survey. Respondents were asked for their age at the birth of their first child. The Mainline Protestant category includes Methodist, Presbyterian, and Lutheran women. Information was not available on Jehovah's Witnesses or Pentecostals.

the Catholic dummy variables remains unchanged over Table 8. Baptist women are significantly less likely to work in jobs with a large number of usual hours worked per week. Jewish women, on the other hand, are likely to work in careers that require more hours worked per week. In spite of controlling for children and education, channels through which differences in religious ideology are likely to manifest themselves, there are still significant differences in career flexibility by religion.

More highly educated women are generally older when their first child is born. If women in certain religious denominations postpone childbearing until later in life, I could be underestimating the number of children these women have, and thus the impact of fertility expectations on job flexibility for these women. Table 12 examines the mean age at first birth for women of different denominations. Because the PSID does not ask women for their age at first birth, the data comes from the 1998 General Social Survey. Information was not available on whether a woman was Episcopalian, Pentecostal, or Jehovah's Witness. Because of this, the Mainline Protestant category includes Methodist, Presbyterian, and Lutheran women only.

The average age at first birth for Baptist women is 21.5, which is significantly younger than women of any other denomination. Mainline Protestant women have their first child at the average age of 23.2 and Catholic women at 23.6. Jewish women have their first child at 26 on average, significantly later than women of any other religious denomination.

Because childbearing appears to occur at different times for different denominations, I estimate equation 2.2 for women between the ages of 30 and 45 only as an additional test of robustness. By the age of 30, the oldest child of the average Baptist woman will be 9 years old and the average Jewish woman's oldest child will be 4. Since Baptist women have more children on average than Jewish women, both are likely to still have children in the household at the age of 45. Table 13 shows the results of this robustness check. The results are similar to those in Table 8. The loss of significance in two of the four specifications for Jehovah's Witnesses and Pentecostal women is likely due to the small sample; approximately 50% of the original sample falls outside of this age range. In this restricted age range, Jewish women are more likely to work from home and work part time than women with no religious preference.

G. Conclusion

Women are found in systematically different careers than are men. Past research addressing whether this difference in career types is optimal choice or discrimination has met with mixed results. One of the issues that makes this problem difficult is the importance of expectations in the career choice process. If a woman expects to have a large family or many in-home responsibilities, she may be more likely to choose a career that is compatible with this choice. Expectations are difficult to measure

Table 13. Religion as a Predictor of Career Flexibility for Women between the Ages of 30 and 45.

	Flexibility Measure			
	Fraction part time	Fraction work from home	Average usual hours worked	Depreciation
Jehovah's Witness	0.116*	0.107	-1.874**	-0.005
	(0.06)	(0.06)	(0.59)	(0.00)
Pentecostal	0.071	0.066	-0.974**	-0.001
	(0.04)	(0.04)	(0.37)	(0.00)
Catholic	0.101***	0.099***	-0.076	-0.000
	(0.02)	(0.02)	(0.22)	(0.00)
Baptist	0.020	0.016	-0.768***	-0.000
	(0.02)	(0.02)	(0.21)	(0.00)
Mainline Protestant	0.004	0.002	-0.040	-0.001
	(0.02)	(0.02)	(0.22)	(0.00)
Jewish	0.110*	0.111*	1.685***	0.006**
	(0.05)	(0.05)	(0.49)	(0.00)
Constant	0.205***	0.194***	43.191***	0.048***
	(0.02)	(0.02)	(0.19)	(0.00)
Observations	5046	5046	3960	3960
F-statistic	9.370	9.218	11.455	2.228

Notes: *Indicates significance at the 10% level, ** at the 5%, *** at the 1% level. Standard errors in parenthesis. Average usual hours worked, fraction who work from home, and fraction part time are all defined for each of the 505 Census occupation codes from the 1990 1% sample using men between the ages of 25 and 55. Depreciation is defined for 17 occupational categories as described in the text. The individual level data is from the PSID for women between the ages 30 and 45.

and are not always fulfilled and ignoring their importance can lead one to incorrect conclusions about the cause of a woman's career path.

Different religious groups have different beliefs on the importance of child bearing and the division of labor within the home. Because of this, fertility expectations should differ by religious group. If this is the case, including a woman's religion in regressions on job attributes should control for fertility expectations and expected effort expended in household chores. In general, it appears that Jehovah's Witnesses are likely to choose the most flexible careers followed by Pentecostal, Catholic, and Baptist women. There is no significant difference in the flexibility of Protestant groups which are more likely to stress gender equality and women with no religious preference. When Jewish women differ from women with no religious preference, they tend to be in less flexible careers. These results are consistent with the human capital notion that women are choosing different careers than men rather than being forced into different job paths.

CHAPTER III

HOUSEHOLD SPECIALIZATION AND THE MALE MARRIAGE WAGE PREMIUM

A. Introduction

Empirical results show that married men earn between 10 and 30% higher wages than their single counterparts, and this wage gap has not been fully explained. The primary explanations for this phenomenon are employer discrimination towards married men, differential selection of men into marriage, or increased productivity as a result of greater specialization of labor for married men. Most studies focus on differentiating between the selection and productivity hypotheses, and the existing literature provides mixed results as to which of these factors is responsible for the wage premium.

It is possible that the marriage wage premium has not been fully explained because the key variables, productivity, household specialization, and ability, are very difficult to define empirically. Many previous studies (Korenman and Neumark 1991, Cornwell and Rupert 1997, Gray 1997, and Hersch and Stratton 2000) control for unobservable individual factors such as ability by using fixed effects models. Most of these studies find a marriage premium of between 5 and 10% even after controlling for individual-specific fixed effects.

Fixed effects estimation is a good solution to the problem of differential selection into marriage. It is arguably more difficult to find a good control for the degree of specialization within the home. Previous studies have used a wife's number of hours worked as a proxy for specialization. Conflicting results Loh (1996) and Gray (1997) suggest that hours worked is a weak proxy. Even theoretically, the relationship between a wife's market work and her husband's household production is ambiguous. One must consider the income and substitution effects associated with a woman beginning to work. The income effect says that married men with working wives may spend less time on household production because household income is greater. On the other hand, the working wife's time is more valuable, so her husband may spend more time on housework than a man whose wife does not work. Using the employment status of the wife as a proxy for specialization implies that the substitution effect dominates the income effect; research by South and Spitze (1994) indicates that this is not the case.

The theoretical uncertainty in the relationship between wife's employment and household production suggests that the number of hours worked by the wife may not be the best proxy for specialization. Hersch and Stratton (2000) use self-reported information on time spent by men in nine different household production activities as a measure of household specialization. They find that time spent on housework has a negative effect on wages, but has no effect on the magnitude of the marriage premium. Their puzzling conclusion is that neither specialization nor selection can explain the wage premium. Self-reported assessment of time spent on household tasks may not fully capture the dynamic of household specialization, and they cannot address the discrimination hypothesis.

Despite the mixed results, increased productivity through specialization of labor is still the prevailing explanation for the marriage wage premium. This paper provides an important contribution to the literature by presenting an alternative measure of specialization within the home, religious denomination. For most men, religion is arguably exogenous. Most Americans adopt the religion of the families they were born into. In addition, certain religious denominations place an emphasis on traditional gender roles. The statement of faith of the Southern Baptist church, for example, says that

> A wife is to submit herself graciously to the servant leadership of her husband even as the church willingly submits to the headship of Christ. She, being in the image of God as is her husband and thus equal to him, has the God-given responsibility to respect her husband and to serve as his helper in managing the household and nurturing the next generation[1].

In a study of religious denomination and gender attitudes, Brinkerhoff and MacKie (1984) investigate differences by denomination in agreement to statements like "when a husband and wife both work, housework should be shared equally." They find that Pentecostals are dramatically less egalitarian than the other denominations studied. Baptists and Catholics have more defined gender roles than those with no religious preference.

In this paper, I investigate whether specialization of labor is higher in families whose religious denomination emphasizes traditional gender roles. I will then examine whether the marriage premium is higher for men of these religious groups. If men in more traditional religious denominations enjoy a higher marriage wage premium, this is evidence that household specialization of labor is an important cause of the wage premium. Unless the process of selection into marriage differs systematically with religious denomination, the selection hypothesis cannot explain differing returns to marriage along denominational lines. This approach will also provide a means to test the discrimination hypothesis, unless employers prefer married men of certain religious denominations over others.

[1] http://www.sbc.net/bfm/bfm2000.asp

B. Literature Review

Hill (1979) was among the first to thoroughly investigate the marriage wage premium. She uses the PSID and examines the wage effects of marital status for white men after carefully controlling for work experience, training, and labor force attachment. The marriage premium remains around 30% even after these controls for worker qualifications are introduced. This suggests that marital status is not a proxy for differential work experience, on-the-job training, or labor force attachment and is instead due to increased productivity through marriage.

Korenman and Neumark (1991) also examine whether marriage makes men more productive. They use the National Longitudinal Study of Young Men as well as company personnel data. The NLSYM provides evidence against the selection hypothesis. In the cross-sectional analysis, the marriage premium is 16.6%. After fixed effects estimation, the marriage premium is 14.8%. Ninety-two percent of the premium remains after controlling for individual fixed-effects. The personnel data reveals that married men are found in higher job grades than single men and that this is the cause of their higher wages. Married men are also more likely to be given high performance ratings which increase the probability of promotion. The findings from the personnel data could be consistent with employer discrimination, but Korenman and Neumark do not investigate this possibility.

Cornwell and Rupert (1997) perform the same analysis as in Korenman and Neumark but extend the panel from the National Longitudinal Survey of Young Men by 5 years and include job tenure as an explanatory variable. These relatively minor changes lead to a completely different conclusion as to the source of the marriage wage premium; they find that the marriage premium from the fixed effects model is no more than 5% to 7%. They also construct a “to-be-married” variable and find that men who are not married in 1971 but marry sometime during the sample earn at least as much as those who are already married. This, too, is evidence against the productivity hypothesis. Their sample size is small (666 men) and confined to men who were between the ages of 19 and 29 in 1971 and therefore may not generalize to other cohorts of men.

Loh (1996) also uses a small age range and concludes that the marriage premium is not due to differences in productivity. His sample includes men in the National Longitudinal Survey of Youth Labor Market Experience who were between the ages of 14 and 22 in 1990. He uses the wife's labor force participation as a proxy for specialization within the home and finds that the marriage premium does not diminish when this control is added. He also examines the self-employed; if marriage makes men more productive in general, self-employed married men should make more money than self-employed single men. Loh finds that self-employed married men earn less than their single counterparts which is clearly contrary to the productivity hypothesis. The productivity hypothesis also suggests that men who co-habit prior to marriage should earn a higher wage premium than those who did not. This is because partners who live together prior to marriage gain information on the other's strengths and weaknesses and can use this information to specialize more efficiently. Currently married men who lived with their wives before marriage receive the same premium as married men who did not live with their spouse prior to marriage. Loh takes this as evidence against the specialization hypothesis.

If the marriage premium is due to selection of higher ability men into marriage, marriage and wages are jointly endogenous. Nakosteen and Zimmer (1987) use instrumental variables estimation in the 1977 PSID to control for this possible endogeneity. They use parent's educational attainment, number of siblings, and dummy variables for race, urban upbringing, whether the respondent was Catholic, and the presence of older siblings as instruments for marriage. The marriage wage premium in the IV estimation is unchanged over the OLS estimate of around 45%, but becomes insignificant due to a large increase in standard error. Nakosteen and Zimmer attribute this loss of significance to evidence for the selection hypothesis. Subsequent researchers have questioned this result because of the difficulty in finding valid instruments for marriage and the imprecision of their findings.

It is possible that the conflicting results from the above studies are due to the changing nature of the marriage premium over time. The marriage wage premium fell by more than 40 percent in the 1980s. Between 1976 and 1980, married men earned 11% more than never-married men. By the 1989-1993 period, married men earned only 6% more than their single counterparts. Gray (1997) uses National Longitudinal Survey of Young Men data from 1976, 1978, and 1980 and National Longitudinal Survey of Youth data from 1989, 1991, and 1993 to examine the changing marriage wage premium. In the early period, the marriage premium appears to be mainly a result of increased productivity of married men. In the 1989-1993 period, however, the fixed effects regressions show no marriage wage premium, evidence that the wage gap during this time period is attributable to selection. These results suggests that the productivity effects of marriage have declined.

This decline could be due to decreased specialization or to diminished returns to specialization within the home. In order to shed light on this issue, Gray adds the wife's labor market hours to the regression as a proxy for the degree of specialization. Because the wife's labor market hours are arguably endogenous to her husband's wages, he uses instrumental variables estimation. Conditional on appropriate instruments, Gray finds that the "decline in the productivity effects of marriage results from less specialization taking place in marriages rather than any decrease in the return to specialization." (pg 502).

Hersch and Stratton (2000) put a unique spin on the problem. They use panel data from the National Survey of Families and Households which includes information on time spent in nine different household production activities. They find that time spent on housework has a negative effect on wages, but has no effect on the magnitude of the marriage premium. Including fixed effects lessens the marriage premium from 11% to 9.4%, indicating that selection matters, but explains little of the wage gap. Their results mirror the mixed messages received from the rest of the literature. The marriage premium is not primarily due to selection. Marriage appears to have made these men more productive, but not through household specialization. They suggest that the premium could be due to preferential treatment from employers or because of changes in men's behavior because of greater stability due to marriage. We still have much to learn about the effect of marriage on wages.

C. Data

I use data from the 1991 to 2001 waves of the Panel Study of Income Dynamics (PSID). The PSID is a longitudinal data set collected by the University of Michigan. This is an ideal data set for this analysis because it has been used extensively in the literature on the male marriage premium, includes questions on a respondent's religious preference, and allows me to identify many different religious denominations. The denominations I consider are Pentecostal, Catholic, Baptist, Jewish, and Mainline Protestant, which includes Presbyterian, Methodist, Episcopalian, Lutheran, and "Other Protestant." The reference category is individuals whose religious preference is "none, atheist, agnostic." There are 686 individuals who do not fall into these categories and who have been omitted from the sample. Most of these men did not answer the religion question or responded that their religious preference was "other".

Individuals who fell into the category of "Other non-Christian: Muslim, Rastafarian, etc" and those who simply responded that they were "Christian" were also omitted.

Because wage and religious preference information is only available for heads of household and their wives, the sample is restricted to men between the ages of 25 and 55 who are the head of the household. The PSID is set up such that if an adult male is present in the household, he is considered the head. I restrict the sample to include only men between the ages of 25 and 55. Men younger than 25 are not included because of the possibility that they are combining school and work, which would likely affect wages. I also exclude men who are nearing retirement. Marriage histories are complete only for men who have been married less than 3 times, so the sample is limited to this group. Less than 5% of the sample is affected by this restriction and results are robust to their inclusion. The marriage variable is defined as men who are legally married at the time of the interview. The PSID occasionally asks whether men are cohabitating at the time of the survey. Because this information is not available in all years, it is not considered in this paper. Because religious groups differ in their stance on cohabitation prior to marriage, interpretation of this variable would be difficult.

Wages are defined as total annual labor earnings of the respondent divided by his annual hours of work on all jobs, including overtime. Wage data was top and bottom coded in each year. Wages that fell below the second percentile were assumed to be equal to the wage level at the second percentile. Wages that were above the ninety-eighth percentile were set equal to the wage level at the 98^{th} percentile. Men who did not work in a given year are excluded from the analysis. The natural log of wages is used in all regression specifications. Experience is defined as age minus years of education minus 6.

D. Empirical Results

Table 14 gives summary statistics at the time of first entry into the sample. The "Single" category includes never married men as well as those who are divorced or widowed. Both wages and non-wage characteristics differ for single and married men. Married men are 2 years older on average than their non-married counterparts. Because of the age difference, Table 14 includes age-adjusted differences between married and single men. Single men have an hourly wage of $11.90 and married men of $13.73. Married men have approximately 2 more years of experience than single men on average, but this is due almost exclusively to the differences in their ages. Married men also work significantly more hours than single men, 134 hours per year after adjusting for age. Married and single men have very similar years of education on average. Men in the sample who are married have been married for an average of ten and a half years. Approximately 27% of men in the sample are Baptist, 24% are Protestant, 34% are Catholic, 2% are Pentecostal, 2% Jewish, and 10% have no religious preference. There are differences in the composition of religious preference by marital status as well. Catholic men in the sample are significantly more likely to be married while Baptist men and atheist, agnostic, and men with no religious preference are more likely to be single.

Table 15 compares variable means by marital status and religious group. There are significant differences both across religious denomination and by marital status. Jewish men have the highest hourly wage regardless of marital status, but married Jewish men make nearly $1.82 per hour more than single Jewish men. After adjusting for age, married Baptist men have the largest wage advantage of their single counterparts and work 140 hours more than single Baptist men. Married Jewish men, on the other hand, work almost 400 more hours per year than single Jews. There were

Table 14. Summary Statistics at Time of First Entry into Sample

	Married		Single			
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Age-adjusted difference	Std. Err
Age	36.396	8.104	34.198	7.911	-	-
Number of children	1.496	1.269	0.544	1.015	0.958	0.042
Hourly wage	13.726	8.611	11.906	8.229	1.444	0.286
Experience	17.782	8.877	15.658	8.623	-0.146	0.100
Annual hours worked	2190	654.9	2056	758.6	134.3	22.9
Years of Education	12.614	2.998	12.539	2.715	0.146	0.010
Years Married	10.253	8.583	0.000	0.000	8.931	0.204
Hours of housework per week	7.782	7.891	8.173	8.145	-0.399	0.270
Proportion Jewish	0.020	0.139	0.016	0.124	0.003	0.005
Proportion Pentecostal	0.028	0.165	0.017	0.128	0.013	0.005
Proportion atheist/agnostic	0.088	0.284	0.151	0.359	-0.059	0.010
Proportion Protestant	0.250	0.433	0.218	0.413	0.029	0.015
Proportion Catholic	0.355	0.479	0.280	0.449	0.073	0.016
Proportion Baptist	0.259	0.438	0.319	0.466	-0.059	0.015
Observations	4693		1083			

Notes: Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. The wage is defined as the total annual labor income of the respondent divided by annual hours worked.

not large differences in education by marital status for any of the religious groups. Married men do fewer hours of housework on average than single men; Jewish men have the largest average reduction of housework hours per week at approximately 2 hours per week.

Table 15 suggests that housework behavior and the marital wage premium differs by religious denomination. While cross-sectional variation provides interesting information, we would really like to know how an individual's wages and in-home responsibilities change upon marriage. As a baseline specification, I estimate the following equation for the sample as a whole:

$$\ln W_{it} = \beta X_{it} + \delta \text{Married}_{it} + A_i + \varepsilon_{it} \tag{3.1}$$

where $\ln W_{it}$ is the log of the hourly wage for individual i at time t, X_{it} is a vector of observable individual characteristics expected to influence the wage, Married_{it} is a dummy variable indicating whether the individual was legally married in a particular

Table 15. Variable Means by Religious Denomination

Baptist				
Variable	Married	Single	Age-adjusted difference	Std. Err
Hourly Wage	12.464	10.207	2.104	0.453
Experience	17.417	16.173	-0.381	0.130
Annual hours worked	2145.751	2008.462	139.316	43.185
Years of Education	12.605	12.301	0.381	0.130
Hours of housework	7.079	8.430	-1.322	0.485
Catholic				
Variable	Married	Single	Age-adjusted difference	Std. Err
Hourly Wage	13.748	12.832	0.562	0.563
Experience	18.884	15.845	0.064	0.227
Annual hours worked	2194.086	2020.381	176.594	39.553
Years of Education	11.822	12.073	-0.064	0.227
Hours of housework	8.675	8.386	0.250	0.536
Protestant				
Variable	Married	Single	Age-adjusted difference	Std. Err
Hourly Wage	14.467	12.708	1.338	0.579
Experience	17.240	15.898	-0.348	0.170
Annual hours worked	2259.389	2152.994	102.532	50.095
Years of Education	13.506	13.127	0.348	0.170
Hours of housework	7.410	8.291	-0.913	0.515
Jewish				
Variable	Married	Single	Age-adjusted difference	Std. Err
Hourly Wage	23.149	17.437	1.827	3.287
Experience	17.699	6.647	1.011	0.024
Annual hours worked	2261.906	2137.447	395.683	197.479
Years of Education	15.667	15.706	0.078	0.539
Hours of housework	5.815	7.824	-1.979	1.817
Pentecostal				
Variable	Married	Single	Age-adjusted difference	Std. Err
Hourly Wage	11.359	10.822	1.052	1.622
Experience	16.870	18.389	0.535	0.617
Annual hours worked	2083.261	2066.822	2.426	156.467
Years of Education	11.527	11.889	-0.535	0.617
Hours of housework	8.146	7.222	0.965	2.016

Notes: Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. The wage is defined as the total annual labor income of the respondent divided by annual hours worked.

year, and A represents any time-invariant, unobserved individual characteristics such as ability or physical attractiveness. Fixed effects estimation allows for the removal of these time-invariant, unobserved individual characteristics and consistent estimation of the returns to marriage. Because of the age restriction, years of education for an individual in the sample are not likely to change over time. Years of education is not significant in any specification and thus not included in X. All results are robust to its inclusion. Experience is included in X in all specifications as a joined spline. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years. Murphy and Welch (1990) show that returns to experience vary greatly over the span of a career. Career growth for high school graduates in the first 10 years of their career is 54%, growth during the next 15 years is 18%, and the decline in the subsequent 15 years is 5%. Using a typical quadratic in experience understates early career wage growth by almost 50% and understates middle career wage growth by around 30%. The joined spline allows wage growth to differ over these career periods.

Table 16 contains the results of estimating equation 3.1. Married men earn about 6% more than single men after controlling for experience. Consistent with research on age-earnings profiles, the returns to experience are greatest during the first 15 years of a career and decline after this point.

The heart of the productivity hypothesis is that marriage allows for increased specialization of labor which allows men to devote more time and effort to market work, thus earning higher wages. While self-reported hours of housework per week do not fully capture specialization of labor, it should shed some light on whether there are differences by religion in housework responsibilities upon marriage. I estimate the

Table 16. Fixed Effects Estimates of the Marriage Wage Premium

Variable	**Coefficient**	**(Std. Err.)**
Married	0.055	(0.015)
Early career experience	0.076	(0.003)
Mid-career experience	0.038	(0.002)
Late career experience	0.025	(0.005)
Intercept	1.315	(0.036)
N	27575	
$F_{(5787,21787)}$	452.531	

Notes: All coefficients are significant at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. The wage is defined as the natural log of the total annual labor income of the respondent divided by annual hours worked. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

following equation:

$$\text{Housework}_{it} = \beta X_{it} + \delta \text{Married}_{it} * \text{Denomination}_i + A_i + \varepsilon_{it}. \tag{3.2}$$

where Housework is defined in two ways: the weekly hours of housework done by the head of the household and the fraction of housework done by the head of household. Fraction of housework done by head of household is included because total household production likely changes with marriage. If home production is positively related to the number of hours spent on housework, a man might increase his housework by 1 hour per week upon marriage but benefit from a 10 hour increase in time allocated to total household production. Simply looking at the number of hours of housework he does per week will not capture this, but the percentage of housework he does will. Both of these could arguably affect wages. The X vector includes controls for number of children in the household and number of years married. I include a control for the number of children in the household because fertility differs across religious groups, and time spent in child care is included in reported hours of housework. Years married is included because specialization within the home is expected to increase over the life of a marriage.

Table 17 contains the results of the estimation of equation 3.2. Men with no religious preference do significantly more hours of housework after marriage, but do a significantly smaller fraction of total housework. This pattern is true for Pentecostal men as well; their fraction of total housework decreases significantly, but they do (insignificantly) more hours of housework after marriage. Baptist, Catholic, and Protestant men do significantly fewer hours of housework and a smaller fraction of total housework. As expected, total hours of housework increases when children are added to the family. This increase falls primarily on the mother, however, as fraction of housework done by the head of household decreases with additional children.

Table 17. Fixed Effects Regressions of Housework Done by Head of Household

	Hours of housework of head of household	**Fraction of housework of head of household**
Variable	**Coefficient** (Std. Err.)	**Coefficient** (Std. Err.)
Married	1.362**	-0.420***
	(0.51)	(0.01)
Married*Baptist	-1.758**	-0.047**
	(0.61)	(0.02)
Married*Catholic	-1.742**	-0.044**
	(0.62)	(0.02)
Married*Protestant	-1.391*	-0.093***
	(0.61)	(0.02)
Married*Jewish	-1.718	0.032
	(1.19)	(0.03)
Married*Pentecostal	1.451	-0.125*
	(1.95)	(0.05)
Number of Children	0.263***	-0.040***
	(0.06)	(0.00)
Years married	-0.083***	-0.004***
	(0.01)	(0.00)
Intercept	8.073***	0.850***
	(0.19)	(0.01)
N	27369	26972
F-statistic	8.153	1531.210

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55.

Controls are also included for number of years married. The longer a couple has been married, the less time the husband spends on housework.

If changes in housework responsibilities capture some dimension of specialization of labor, Table 17 suggests that Protestant, Catholic, and Baptist men should enjoy a significant marriage premium. It is unclear what to expect from Jewish and Pentecostal men. Brinkerhoff and MacKie's (1984) findings suggest that Pentecostal men will benefit most from specialization of labor, followed by Baptists, Catholics, and Protestants. To investigate whether the marriage wage premium differs by religious denomination, I estimate the following equation:

$$\ln W_{it} = \beta X_{it} + \delta \text{Married}_{it} * \text{Denomination}_i + A_i + \varepsilon_{it}. \tag{3.3}$$

Fixed effects estimates of equation 3.3 can be found in Table 18. Men with no religious preference do not experience an increase in wages upon marriage while Jewish, Catholic, Baptist, Protestant, and Pentecostal men do. In comparison to the 6% marriage wage premium in Table 16, these premiums are very large. With 90% confidence, Baptist men receive a wage premium over men with no religious preference of between 0.61% and 19%, Catholic men between 3% and 23%, Protestant men 5.4% to 25%, and Jewish men 3.3% to 45%. Pentecostal men receive the largest wage premium with a range of 37% to an incredible 138%. This is consistent with Brinkerhoff and MacKie's finding that Pentecostal families are the least egalitarian.

It is possible that the full benefits of specialization of labor within the household are not realized the moment one gets married. Korenman and Neumark (1991) find that the effects of marriage appear gradually over time rather than from a one time intercept shift upon marriage. If couples get better at specializing in the household over time, we would expect a significant growth in marriage premium with number of years married. In Table 19, I interact years married with religious denomination so that wage growth can differ by religion. There is still a significant jump in wages at marriage for Baptist, Catholic, Protestant, and Pentecostal men; the point estimates are similar to those in Table 18 but are higher for all groups but Catholic men. These men appear to benefit from marriage immediately, as opposed to gaining as the marriage ages. There is no longer an intercept shift for Jewish men; however, they seem to benefit from marriage in the form of a larger growth rate, approximately 2% per year married. The average Jewish man in the sample has been married for 12.7 years which implies a wage premium of 24%. The wage premium may diminish over

Table 18. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination

Variable	**Coefficient**	**(Std. Err.)**
Married	-0.051	(0.043)
Married*Baptist	0.091*	(0.051)
Married*Catholic	0.117**	(0.052)
Married*Protestant	0.137***	(0.051)
Married*Jewish	0.201**	(0.102)
Married*Pentecostal	0.591***	(0.167)
Early career experience	0.076***	(0.003)
Mid-career experience	0.038***	(0.002)
Late career experience	0.024***	(0.005)
Intercept	1.305***	(0.036)
N	27575	
$F_{(5792,21782)}$	203.222	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

time for Baptist men.

The argument for increased wage growth as a marriage progresses is that over time partners gain information on the other's strengths and weaknesses and can use this information to specialize more efficiently. Perhaps rigidly defined gender roles in certain denominations preclude these gains over time. If there are denominational norms in the kind of jobs that a wife takes over upon marriage and very little deviation from these norms, most if not all of the benefit of marriage would be realized at the moment the marriage begins. Perhaps Pentecostal women, for example, immediately begin cooking meals, doing laundry, and washing dishes upon marriage because that is a cultural norm in the denomination. There are not additional benefits over time because traditional "women's work" does not change. It may be that if the woman is more well suited to changing the oil in the car than her husband, the household never benefits from this because she never takes on this responsibility.

E. Tests of Robustness

Table 20 examines the changes in the magnitude of the wage premium for different religious groups when additional control variables are added. The specification in the first column of Table 20 is the same as in Table 19, but controls for 9 occupational categories are added. If men of different religious groups have differing tastes or access to certain careers, failing to control for occupation will lead to biased estimates. Household service occupations are the omitted category. Results are very similar to those in Table 19. Men with no religious preference continue to experience no increase in wage levels or wage growth with marriage. The large premium for Pentecostal men persists; the wage premium they earn is 84% higher than that of men with no religious preference after controlling for observable characteristics. There remains a significant

Table 19. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination: Premium Allowed to Vary with Years Married

Variable	**Coefficient**	**(Std. Err.)**
Married	-0.068	(0.047)
Married*Baptist	0.121**	(0.056)
Married*Catholic	0.099*	(0.057)
Married*Protestant	0.140**	(0.056)
Married*Jewish	0.091	(0.109)
Married*Pentecostal	0.597***	(0.169)
Years married	0.005	(0.004)
Years married*Baptist	-0.009*	(0.005)
Years married*Catholic	0.001	(0.005)
Years married*Protestant	-0.002	(0.005)
Years married*Jewish	0.021**	(0.008)
Years married*Pentecostal	0.004	(0.010)
Early career experience	0.074***	(0.003)
Mid-career experience	0.036***	(0.002)
Late career experience	0.024***	(0.005)
Intercept	1.320***	(0.037)

N	27452
$F_{(5771,21680)}$	123.041

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

wage premium for Baptist men of 11.5% and Protestant men of around 13% over men with no religious preference. Jewish men continue to experience a 2% wage growth for each additional year married.

In column 2, I add controls for hours of housework per week and number of children to the specification in column 1. While Hersch and Stratton (2000) find that housework has a negative effect on wages, I do not find a significant effect of hours of housework per week on wages. The addition of children to the family has a positive effect on wages; perhaps because men commit more effort to their jobs with the added responsibility of providing for a family. Coefficient magnitudes are nearly identical to those in column 1. The third column of Table 20 includes controls for experience, experience^2, experience^3, and experience^4 instead of the spline in experience. Again, coefficient estimates change very little and the large wage premium for Pentecostal men persists.

Annual hours of work also clearly differs by religious denomination and has an effect on wages. Table 21 investigates how including annual hours worked affects the coefficient estimates in Table 19 and Table 20. The specification in the first column of Table 21 is the same as in Table 19, but controls for annual hours worked have been added. The inclusion of hours of work has very little effect on the coefficient estimates. Pentecostal men still enjoy a large and significant wage premium. Baptist, Catholic, and Protestant men also experience an intercept shift upon marriage. Married Jewish men do not have a significant intercept shift, but benefit from marriage in terms of higher wage growth. In the second column of Table 21, I control for 9 occupational categories. The magnitude of the estimates is similar to those in column 1. The large premium for Pentecostal men persists; they earn an average premium of 86% after controlling for observable characteristics. There remains a significant wage premium for Baptist and Protestant men. Jewish men continue to experience a 2% wage growth

Table 20. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination: Controls for Occupation, Number of Children, Housework, and Additional Experience Added

Variable	Coef	(S.E)	Coef	(S.E.)	Coef	(S.E.)
Married	-0.044	(0.05)	-0.053	(0.05)	-0.062	(0.05)
Married*Baptist	0.109*	(0.06)	0.107*	(0.06)	0.115**	(0.06)
Married*Catholic	0.080	(0.06)	0.078	(0.06)	0.085	(0.06)
Married*Protestant	0.122**	(0.06)	0.118**	(0.06)	0.123**	(0.06)
Married*Jewish	0.058	(0.11)	0.065	(0.11)	0.029	(0.11)
Married*Pentecostal	0.612***	(0.17)	0.600***	(0.17)	0.616***	(0.17)
Years married	0.004	(0.00)	0.004	(0.00)	0.005	(0.00)
Yearsmarried*Baptist	-0.009*	(0.00)	-0.008*	(0.00)	-0.009*	(0.00)
Years married*Catholic	0.001	(0.00)	0.001	(0.00)	0.000	(0.00)
Years married*Protestant	-0.003	(0.00)	-0.002	(0.00)	-0.003	(0.00)
Years married*Jewish	0.022***	(0.01)	0.022***	(0.01)	0.021**	(0.01)
Years married*Pentecostal	-0.005	(0.01)	-0.004	(0.01)	-0.005	(0.01)
Early experience	0.074***	(0.00)	0.073***	(0.00)		
Mid-career experience	0.037***	(0.00)	0.038***	(0.00)		
Late career experience	0.026***	(0.00)	0.028***	(0.00)		
Experience					0.159***	(0.02)
Experience^2					-0.006***	(0.00)
Experience^3					0.000**	(0.00
Experience^4					-0.000*	(0.00)
Professional, Technical	0.095***	(0.03)	0.092***	(0.03)	0.088***	(0.03)
Managers	0.084***	(0.03)	0.082**	(0.03)	0.080**	(0.03)
Sales	0.050	(0.03)	0.050	(0.03)	0.047	(0.03)
Clerical	0.046	(0.03)	0.042	(0.03)	0.044	(0.03)
Craftsmen	0.093***	(0.02)	0.092***	(0.02)	0.091***	(0.02)
Operatives	0.090***	(0.03)	0.089**	(0.03)	0.087**	(0.03)
Transport Operatives	0.070**	(0.03)	0.062**	(0.03)	0.062**	(0.03)
Laborers	0.074***	(0.03)	0.072***	(0.03)	0.070**	(0.03)
Farmers	-0.056	(0.06)	-0.057	(0.06)	-0.061	(0.06)
Number of children			0.015**	(0.01)	0.010*	(0.01)
Weekly housework hours			0.001	(0.00)	0.001	(0.00)
Intercept	1.252***	(0.04)	1.257***	(0.04)	0.901***	(0.08)
N	26536		26460		26460	
F-statistic	78.813		72.636		71.483	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Annual hours worked has been divided by a thousand. Household service occupations are the omitted occupational category.

for each additional year married. In column 3, controls for hours of housework per week and number of children are added to the specification in column 2. Coefficient magnitudes change very little. Pentecostal men still experience a 83% wage premium at marriage, and Jewish men an average of 2% wage growth per year married. The fourth column of Table 21 includes controls for experience, experience^2, experience^3, and experience^4 instead of the spline in experience. Coefficients change very little.

It is possible that age at first marriage differs by religious denomination. If most men in a particular denomination marry before the age of 25, the wage premium for that denomination will be identified only on the men who marry later than average, which may not be a representative sample. Table 22 shows the mean and median age at first marriage by religious denomination. Pentecostal men marry earliest, with a median age at first marriage of 22.5. Protestant and Baptist men are 23 on average at their first marriage, and Catholic men and men with no religious preference are 24. Jewish men marry the latest, with a median age at first marriage of 25.

Because men of some denominations marry earlier than others, I estimate equation 3.3 for men between the ages of 20 and 55 as an additional test of robustness. This expanded age range includes over 75% of all marriages in the sample. Because of the possibility that some of these men are combining school with work, an indicator variable for being currently enrolled in school is included in this specification. The results of this estimation can be found in Table 23 and are very similar to the results in Table 18. Men with no religious preference do not experience an increase in wages upon marriage while Jewish, Catholic, Protestant, and Pentecostal men do. There is no wage premium for Baptist men in the expanded age range. While the coefficient magnitudes are slightly smaller in this specification, there is still an average wage premium over men with no religious preference of 11.63% for Catholic men, 14% for Protestant men, 30.34% for Jewish men, and 47.55% for Pentecostal men.

Table 21. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination: Controls for Annual Hours Worked, Occupation, Number of Children, Housework, and Additional Experience Added

Variable	Coef	(S.E)	Coef	(S.E.)	Coef	(S.E.)	Coef	(S.E.)
Married	-0.063	(0.04)	-0.047	(0.05)	-0.055	(0.05)	-0.065	(0.05)
Married*Baptist	0.112**	(0.05)	0.109**	(0.05)	0.104*	(0.05)	0.113**	(0.05)
Married*Catholic	0.114**	(0.05)	0.090	(0.06)	0.087	(0.06)	0.094*	(0.06)
Married*Protestant	0.122**	(0.05)	0.108**	(0.05)	0.101*	(0.05)	0.105*	(0.05)
Married*Jewish	0.113	(0.11)	0.081	(0.11)	0.087	(0.11)	0.046	(0.11)
Married*Pentecostal	0.607***	(0.16)	0.619***	(0.16)	0.606***	(0.16)	0.622***	(0.16)
Years married	0.005	(0.00)	0.005	(0.00)	0.004	(0.00)	0.005	(0.00)
Years married*Baptist	-0.008	(0.00)	-0.007	(0.00)	-0.007	(0.00)	-0.007	(0.00)
Years married*Catholic	0.000	(0.00)	0.000	(0.00)	-0.000	(0.00)	-0.001	(0.00)
Years married*Protestant	-0.003	(0.00)	-0.003	(0.00)	-0.003	(0.00)	-0.003	(0.00)
Years married*Jewish	0.020***	(0.01)	0.021***	(0.01)	0.021***	(0.01)	0.020**	(0.01)
Years married*Pentecostal	0.010	(0.01)	0.001	(0.01)	0.002	(0.01)	0.001	(0.01)
Annual hours worked	-0.261***	(0.01)	-0.269***	(0.01)	-0.271***	(0.01)	-0.271***	(0.01)
Early career experience	0.078***	(0.00)	0.078***	(0.00)	0.076***	(0.0 0)		
Mid-career experience	0.037***	(0.00)	0.038***	(0.00)	0.040***	(0.00)		
Late career experience	0.022***	(0.00)	0.024***	(0.00)	0.026***	(0.00)		
Experience							0.171***	(0.02)
Experience^2							-0.006***	(0.00)
Experience^3							0.000***	(0.00)
Experience^4							-0.000**	(0.00)
Professional, Technical			0.110***	(0.03)	0.107***	(0.03)	0.104***	(0.03)
Managers			0.109***	(0.02)	0.107***	(0.02)	0.105***	(0.02)
Sales			0.056*	(0.03)	0.056*	(0.03)	0.053*	(0.03)
Clerical			0.050*	(0.03)	0.047*	(0.03)	0.049*	(0.03)
Craftsmen			0.113***	(0.02)	0.112***	(0.02)	0.111***	(0.02)
Operatives			0.101***	(0.03)	0.099***	(0.03)	0.097***	(0.03)
Transport Operatives			0.084***	(0.03)	0.075***	(0.03)	0.075***	(0.03)
Laborers			0.058**	(0.03)	0.056**	(0.03)	0.054**	(0.03)
Farmers			0.022	(0.05)	0.019	(0.05)	0.016	(0.05)
Number of children					0.017***	(0.01)	0.012**	(0.01)
Weekly housework hours					-0.000	(0.00)	-0.000	(0.00)
Intercept	1.829***	(0.04)	1.782***	(0.04)	1.798***	(0.04)	1.404***	(0.08)
N	27452		26536		26460		26460	
F-statistic	229.144		149.463		138.739		135.840	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Annual hours worked has been divided by a thousand. Household service occupations are the omitted occupational category.

Table 22. Mean and Median Age at First Marriage by Religious Preference

Group	Obs	Median	Mean	Std. Err.	Std. Dev.
Pentecostal	130	22.5	23.900	0.451	5.141
Protestant	1196	23	23.890	0.136	4.696
Baptist	1336	23	23.994	0.146	5.325
Catholic	1841	24	24.460	0.112	4.786
None	551	24	24.895	0.227	5.322
Jewish	91	25	26.077	0.456	4.349

Notes: Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55.

As expected, men who are combining school with work earn lower wages.

Fixed effects estimates do not allow for the consideration of variables that do not vary over time. Because of this, education in fixed effects estimation is only identified from men who get more education while they are in the sample. Men of different religious denominations have different levels of education on average. Wage premiums vary with education, and it is possible that the institution of marriage functions differently for men with different levels of education. To investigate these issues, I estimate equation 3.3 separately for men with less than a college education and men with at least a college education.

Table 24 contains the results of estimating equation 3.3 for men with less than a college education. Men with no religious preference and less than a college education earn significantly *lower* wages after marriage. Married Baptist men with less than a college education earn a 12.19% higher premium upon marriage than men with no religious preference. Pentecostal men in this education category earn an average wage premium of 108%!

Table 23. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination for Men between the Ages of 20 and 55.

Variable	Coefficient	(Std. Err.)
Married	-0.047	(0.038)
Married*Baptist	0.072	(0.046)
Married*Catholic	0.110**	(0.047)
Married*Protestant	0.131***	(0.047)
Married*Jewish	0.265***	(0.097)
Married*Pentecostal	0.389***	(0.130)
Early experience	0.079***	(0.002)
Mid-career experience	0.038***	(0.002)
Late career experience	0.024***	(0.005)
Student	-0.101**	(0.046)
Intercept	1.294***	(0.031)
N	29173	
$F_{(6061,23111)}$	221.98	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 20 and 55. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

Table 24. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination for Men with Less Than a College Education.

Variable	**Coefficient**	**(Std. Err.)**
Married	-0.087*	(0.049)
Married*Baptist	0.115**	(0.058)
Married*Catholic	0.141**	(0.060)
Married*Protestant	0.170***	(0.060)
Married*Jewish	0.023	(0.260)
Married*Pentecostal	0.732***	(0.176)
Early experience	0.066***	(0.003)
Mid-career experience	0.037***	(0.002)
Late career experience	0.027***	(0.005)
Intercept	1.278***	(0.047)
N	20516	
$F_{(4542,15973)}$	116.709	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

The estimates for men with at least a college education are found in Table 25. Interestingly, there is not a significant marriage wage premium by religion for these men. Gains in wages by religious denomination appear for less educated men only. A possible explanation for this phenomenon lies in the difference in physical demands of the jobs of college educated men and non-college educated men. The jobs of less educated men tend to require more physical exertion. Perhaps having the effort saved by having a wife who cooks and cleans at home is more valuable to these men who expend a great deal of physical effort on the job.

F. Conclusion

The cause of the male marriage wage premium is not fully understood. The prevailing hypotheses are differences in productivity due to increased specialization of labor within the home, selection of higher quality men into marriage, or employer discrimination. Differentiating between these hypotheses is difficult in part because a suitable proxy for household specialization is hard to find.

The advantage from marriage that a man will enjoy depends in large part on how many of the household responsibilities his wife is willing to take on. Religious teaching is a powerful means of shaping gender attitudes and expectations of the duties of a husband and wife. Pentecostal men and women, for example, were not as likely to agree that "when a husband and wife both work, housework should be shared equally" as were Catholic husbands and wives (Brinkerhoff and MacKie 1984). Fixed effects estimation shows that Baptist, Catholic, and Protestant men have significant reductions in housework responsibilities after marriage. These facts suggest that if household specialization is the primary cause of the marriage wage premium, men in these denominations should benefit more from marriage than men with no religious

Table 25. Fixed Effects Estimates of Marriage Wage Premiums by Religious Denomination for Men with at Least a College Education.

Variable	**Coefficient**	**(Std. Err.)**
Married	0.082	(0.088)
Married*Baptist	0.006	(0.118)
Married*Catholic	0.021	(0.108)
Married*Protestant	-0.009	(0.103)
Married*Jewish	0.047	(0.134)
Married*Pentecostal	-0.518	(0.526)
Early experience	0.090***	(0.004)
Mid-career experience	0.042***	(0.003)
Late career experience	-0.027	(0.022)
Intercept	1.591***	(0.057)
N	7059	
$F_{(1273,5785)}$	89.154	

Notes: * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. Data is from the 1991 to 2001 waves of the PSID and includes men between the ages of 25 and 55. Returns to experience are allowed to differ between the first 15 years of experience, the second 15 years, and any experience after 30 years.

preference.

When a Pentecostal man marries, his wages increase dramatically, even after controlling for a number of observable characteristics. Catholic men, Protestant men, and Baptist men also benefit from large wage increases after marriage, although the benefit to Baptist men may diminish over time. The wage premium for Jewish men appears to be in the form of increased wage growth. There is no reason to expect that employers would prefer married Pentecostal men to single Pentecostal men more than they prefer married Catholic men to single Catholic men, so discrimination cannot explain these findings. Fixed effects estimation removes time-invariant individual characteristics, and it is difficult to imagine a scenario in which selection into marriage differs across religious denomination by a characteristic that varies over time. These significant differences in marriage premium by religious denomination provide new support for the productivity hypothesis.

CHAPTER IV

ESTIMATING RISK AVERSION AT THE MICRO LEVEL USING LABOR SUPPLY DATA

A. Introduction

Why do some people enjoy bungee-jumping while others do not? Why do some individuals hold extensive insurance portfolios? It is intuitive to attribute these differences between people to their individual preferences for risk, a concept that is appealing but hard to quantify. Kenneth Arrow (1965) and John Pratt's (1964) coefficient of relative risk aversion (RRA) has been the standard risk measure since the mid 1960s. Since that time, theorists and empiricists have studied this measure, each grappling with their own set of problems. Theorists have called expected utility theory into question based on various violations of its predictions and axioms (Allais Paradox, for example), but a dominant alternative theory has not emerged. Many interesting questions about individual risk preferences have not been addressed empirically, likely because the information necessary to compute risk aversion is not available in traditional surveys. Most previous studies of individual risk preferences have been limited to the rare instances where consumption or portfolio data is available. These limitations remain an impediment to progress on topics such as how risk preferences are formed, how they vary in the population, and how they affect decision making and behavior.

Raj Chetty (2006) uses the fact that RRA is directly related to the ratio of the income and compensated wage elasticities of labor supply and presents a formula that allows one to estimate the coefficient of relative risk aversion using labor supply data. His risk aversion measure, γ, is the curvature of utility over wealth. The estimates of γ in his paper are constructed from previous empirical estimates of labor supply elasticities and describe the risk aversion of a representative agent. While the primary purpose of his paper is placing an upper bound on risk aversion, his measure has the potential to be very useful in explaining individual risk preferences since most microeconomic datasets contain the necessary information to estimate his RRA coefficient. In this paper, I estimate Chetty's coefficient of relative risk aversion,γ, at the micro level using the 1996 Panel Study of Income Dynamics (PSID).

Chetty's paper provides new evidence that values of RRA greater than 2 are not consistent with the empirical estimates of the uncompensated wage elasticity of labor supply. Many existing empirical estimates of RRA are greater than 2, however. Chetty contends that this is evidence that the expected utility model does not do a good job of explaining choices under uncertainty in different domains. The 1996 PSID contains a measure of RRA based on an individual's responses to questions on hypothetical gambles. Because of this, I can compare a measure of RRA derived from an individual's labor supply behavior to a measure based on his choices under uncertainty to see if the measure of risk aversion is consistent across these domains. I then use data on health insurance coverage to determine whether each measure correctly identifies individuals with high levels of risk aversion since individuals who insure should be more risk averse than those who do not.

B. Background

1. Measuring Risk Aversion

Chetty (2006) provides the intuition behind estimating γ with labor supply data. Recall that RRA is proportional to $\frac{u_{cc}}{u_c}$ where u_c is the first partial derivative of the utility function with respect to consumption and u_{cc} is the second derivative. An agent's compensated wage elasticity of labor supply is directly related to u_c; the larger the marginal utility of consumption, the more benefit the agent gets from an additional dollar of income, so the more willing he is to work when his wage increases. Income elasticity of labor supply is directly related to u_{cc}; as u_{cc} increases, the agent's marginal utility of consumption decreases significantly as income rises, causing him to work less. Chetty is not the only one to link labor supply choices to risk preferences. Smith et al (2003) use a semi-log specification of the hours of work equation, job risk data, and wage rates to estimate the value of a statistical life.

Using the links between labor supply and the utility function, Chetty gives the following expression for the coefficient of relative risk aversion, γ :

$$\gamma = -(1 + \frac{wl}{y})\frac{\varepsilon_{l,y}}{\varepsilon_{l^c,w}}(y, w) + (1 + \frac{y}{wl})\varepsilon_{u_c,l}, \tag{4.1}$$

where γ is "the curvature of utility over wealth – the parameter that determines risk preferences over immediately-resolved wealth gambles in an expected utility model – when total labor supply l is fixed" (2006, p. 4), w is the wage rate, y is unearned income, l is labor supply, $\varepsilon_{l,y}$ denotes income elasticity of labor supply, $\varepsilon_{l^c,w}$ denotes compensated wage elasticity of labor supply, and $\varepsilon_{u_c,l}$ is the elasticity of the marginal utility of consumption with respect to labor.

Using equation 4.1, Chetty solves for γ in terms of magnitudes that can be estimated empirically (2006, p. 9).

$$\gamma = (1 + \frac{wl}{y})\frac{-\varepsilon_{l,y}}{\varepsilon_{l^c,w}}/(1 - (1 + \frac{y}{wl})[\lim_{\triangle l \to 0} \frac{\triangle c}{c}/\frac{\triangle l}{l}]) \tag{4.2}$$

The numerator of the equation can be easily estimated from a wide variety of data sources. Estimation of $\lim_{\triangle l \to 0} \frac{\triangle c}{c}/\frac{\triangle l}{l}$, on the other hand, requires information on the "consumption choices of agents who face small, permanent exogenous shocks to labor supply" (Chetty 2006, p. 10). This requirement sharply narrows the data from which γ can be estimated. A natural way to estimate this parameter is to look at the changes in an individual's consumption due to job loss or disability. Even this strategy has its problems; estimation of the complementarity parameter requires small fluctuations in l as opposed to the large ones that would come with job loss. Also, job loss is likely to be a temporary fluctuation, and the model requires information about permanent changes. Chetty cites many studies that estimate this parameter in a variety of ways using many different data sources and concludes that $\frac{\triangle c}{c}/\frac{\triangle l}{l} < 0.15$ and is probably closer to 0.1.

Because it is difficult to find the data necessary to satisfactorily estimate the complementarity parameter, and it has been found empirically to be small, I will assume that $u_{cl} = 0$ in order to estimate γ. There is a large body of research that contradicts the assumption that $u_{cl} = 0$, for example, Browning and Meghir (1991) and Browning and Crossley (2001). Ziliak and Kniesner find that consumption and leisure are substitutes and "omitting consumption imparts a downward bias on the non-labor income elasticity of labor supply and an upward bias on the compensated wage elasticity of labor supply" (2005, p. 25).

If $u_{cl} > 0$, my estimates of γ will understate the true γ. To preview future results, with the assumption of additive separability, the mean of γ in my sample is 0.73. If I assume that $\lim_{\triangle l \to 0} \frac{\triangle c}{c}/\frac{\triangle l}{l} = 0.1$, γ for the mean individual increases to 0.82. Unless u_{cl} varies with γ systematically in the population, mis-specifying u_{cl} will cause a mis-measurement in the level of curvature of the utility function, but not the difference in curvature across individuals, which is of primary interest. While not ideal, the assumption of $u_{cl} = 0$ is critical to easily obtaining estimates of γ from a wide variety of data sources. I examine whether this measure captures risk preferences in spite of the limitations imposed by the simplifying assumptions.

The primary focus of Chetty's paper is bounding the coefficient of relative risk aversion. Using previous empirical estimates of labor supply elasticities, he finds γ to have a mean of 1 and an upper bound of 2. Existing empirical estimates of RRA often find much higher values of relative risk aversion. I calculate γ at the individual level and find it to have a mean of 0.733 in the sample. I compare these estimates to an alternative measure of RRA based on hypothetical gambles. This gives me the unique opportunity to compare two different measures of relative risk aversion over the same sample.

2. Labor Supply Estimation

In order to obtain estimates of relative risk aversion, it is necessary to address several issues involved in estimating labor supply. I must decide how wages and non-labor income should be measured, what functional form the labor supply equation will take, and what variables to include as controls. I must also address the life-cycle aspect of the labor supply decision as well as potential endogeneity of wages and non-labor income.

Because different economists take different approaches to the above problems, estimates of wage elasticities in the past 25 years have varied widely. Pencavel (2002) observed that early studies generally estimated static labor supply models and found uncompensated wage elasticities that were small or even negative. More recent studies have largely moved away from static specifications and have found a positive relationship between wages and hours worked. Blundell and MaCurdy (1999) show that if labor supply decisions have life-cycle aspects which are ignored, estimates based on a static model often confuse the effects of movements along the wage profile with shifts in the profile. Individuals plan for retirement, invest in human capital, and engage in other activities that can only be understood in the life-cycle setting. Ignoring these aspects leads to estimated parameters without clear economic interpretation.

Fortunately, there is a cross-section specification that is consistent with life-cycle considerations. Blundell and MaCurdy show that regressing hours of work on non-labor income, age, age-squared, and log-wage allows for calculation of an elasticity that can be interpreted as the within-period effect of a change in wage on work hours under certain assumptions[1] (1999, p. 1603).

Acknowledging life-cycle effects allows for elasticity estimates with economic interpretation, but the estimates must be consistent to be meaningful. There are likely endogeneity problems in labor supply estimation, i.e., variables that affect both wages and hours worked that are unobserved to the researcher. If tastes for work are correlated with wages for example, cross-sectional estimation will provide inconsistent estimates of wage elasticity. Two stage least squares (2SLS) is most often used to overcome the endogeneity problem, but identifying good instrumental variables is an on-going problem in the labor supply literature. In general, age and education in higher order polynomials are used. Most studies include a variety of other instruments that could possibly be correlated with wages without affecting hours of work. These "kitchen sink" instruments are potentially problematic. If correlation between the instruments and the regressor is weak, the distribution of the 2SLS estimator is biased in the same direction as the OLS estimator in finite samples and standard confidence intervals are unreliable (Staiger & Stock 1997). Lee (2001) suggests that in order to limit the bias of the 2SLS estimator, strong instruments must be used. When instrumenting wages, he finds it beneficial to remove all instruments but those that both theoretical and empirical literature have agreed on: years of education and experience.

[1] It must be assumed that the consumer knows that he will work for T periods, that the effects of interest rates or time preferences can be captured by the intercept and other parameters, and that the coefficient on age and age-squared for the lifetime wage and income paths are constant across consumers.

Non-labor income also poses endogeneity problems. People who are less averse to work can save more money early in life that can be turned into non-labor income in the future (Pencavel 2002). Married people may have more incentive to save than their single counterparts because of present and future familial responsibilities; marital status will be used as an instrument for non-labor income. I will analyze the sensitivity of estimates to the instruments chosen in a later section.

3. Previous Estimates of RRA

As noted earlier, empirical estimates of RRA generally require very specific data. Friend and Blume (1975) present some of the earliest empirical estimates. They used 1962 and 1963 Federal Reserve Board data to study the demand for risky assets and concluded that relative risk aversion generally exceeds one and is likely greater than two. Weber (1975) used consumer expenditure data from 1930-1970 and estimated that RRA was between 1.3 and 1.8. Hansen and Singleton (1982) use consumption and stock return data and estimate RRA to lie between .35 and 1. Using the hypothetical gambles questions in the Health and Retirement Study, Barsky et al. (1997) find mean estimates of RRA to be between and 0.7 and 15.8. In their paper on a nonparametric measure of value-at-risk, Aït-Sahalia and Lo (2000), construct an estimator of the coefficient of relative risk aversion in a representative agent equilibrium model. They use the S&P 500 and find that constant relative risk aversion ranges between 1 and 60 with a weighted average of 12.7. Using the Consumer Expenditure Survey, Gourinchas and Parker (2002) estimate a structural model of the life-cycle consumption spending of households. They find the relative risk aversion of an average household to be between 0.5 and 1.4. Halek and Eisenhauer (2001) use life insurance data in the Health and Retirement Study to obtain estimates of RRA that range from 0.029 to 680.83. As their estimates are highly skewed, it is instructive to note that the median value of their measure of RRA is 0.888, and the interquartile range runs from 0.54 to 1.83. Many interesting questions about risk preferences remain unanswered because of the limitations of the above datasets.

C. Data

I use data from the University of Michigan's Panel Study of Income Dynamics for the year 1996. The PSID is a good data set for this analysis because it includes all of the necessary variables on labor supply and has been used extensively in the labor supply literature. The 1996 PSID also included the 'Estimating Risk Tolerance' supplement, whereby respondents' coefficients of relative risk aversion were calculated from their answers to a series of questions like the following:

> Suppose you had a job that guaranteed you income for life equal to your current, total income. And that job was (your/your family's) only source of income. Then you are given the opportunity to take a new, and equally good, job with a 50-50 chance that it will double your income and spending power. But there is a 50-50 chance that it will cut your income and spending power by a third. Would you take the new job?

If the respondent said yes, the cut in income was increased to half. If he would risk half of his income, he is asked if he would accept a potential cut in spending power of 75%. If he would not take the initial gamble, the cut in income is decreased to 20%. If he refuses the 20% income gamble, the proposed cut in income is reduced to ten percent[2]. Based on the individual's response, he is grouped into 1 of 6 categories which can be ranked by risk aversion without any assumptions as to the functional

[2]For more information visit:
http://psidonline.isr.umich.edu/data/Documentation/Cbks/Supp/rt.html

form of the utility function. Respondents who rejected all gambles are placed into the most risk averse category, followed by those who rejected the initial gamble and the 20% income gamble but accepted the gamble of 10%, and so on. The PSID includes the individual's responses to each of the 5 proposed gambles as well as a measure of relative risk tolerance based on these responses.

The PSID risk supplement is based on similar questions in the Health and Retirement Study and the corresponding paper by Barsky et al. (1997). A measure of relative risk tolerance is included in the 1996 data that is calculated following the method described in Barsky et al. (1997). In particular, if one is willing to assume that relative risk aversion is constant over the relevant region, a respondent's answer to the above question places numerical bounds on his coefficient of relative risk aversion. An expected utility maximizer will weigh the expected utility of the gamble against the utility he derives from his current lifetime income. More specifically, he will take the 50-50 gamble of doubling lifetime income at the risk of having it fall by the fraction $1 - \lambda$ if

$$\frac{1}{2}u(2c) + \frac{1}{2}u(\lambda c) \geq u(c),$$

where u is the utility function and c is permanent consumption (Barsky et al. 1997, p. 540).

Since the survey response is likely a noisy measure of a respondent's true risk preference, Barsky et al. (1997) assume that true relative risk tolerance, x, is normally distributed and $x = ln(RRA^{-1})$. The observed relative risk tolerance x^* is in a given risk category if $x \in B_i$ where B_i is the range of log risk tolerance for the categories described above. The utility function is assumed to be $u(c) = \frac{1}{1-RRA}c^{1-RRA}$. A likelihood function is constructed and RRA^{-1} is recovered by computing the expected e^x conditional on being in a particular group (Barsky et al. 1997, p. 545-546, PSID documentation). The PSID data includes this measure of relative risk tolerance, and RRA is the reciprocal of this measure. There are four values of relative risk aversion in the PSID sample, 1.754, 2.857, 3.571, and 6.667. The mean in the sample is 4.584.

For analysis in this paper, the population is defined as male heads of household who are between the ages of 25 and 61. The PSID is set up such that if an adult male is present in the household, he is considered the head. Female heads of household are not considered at this time because there is some concern that the factors governing female labor supply are different from those governing male labor supply. The age restriction is intended to exclude individuals who may be combining school with work and individuals nearing retirement. Also dropped were respondents who reported being retired (714 observations), fulltime students (47 obs), permanently disabled (249 obs), full time housekeepers (14), and those in prison (49). Respondents were dropped if they worked fewer than 100 hours in 1995 (81 observations) or if they reported zero earnings (250 observations). Seventy-two observations with wages of less than $2/hour were also dropped. This leaves a sample size of 3900 individuals. One hundred and four individuals did not report their educational attainment, so they are not included in the regressions. The sensitivity to the age and wage constraints will be examined in the Results section.

Hours of work is defined as weeks worked on respondent's main job times hours of work per week. Annual overtime hours and hours of work on a second job were also included. The wage variable is defined as the total labor earnings of the respondent in 1995 divided by his hours of work. Summary statistics for variables used in the regressions can be found in Table 26.

Table 26. Summary Statistics for Variables Included in Regressions

Variable	Obs	Mean	Std. Dev.	Min	Max
Hours worked	3796	2223.134	608.915	104	5000
Wage	3796	18.047	16.78	2	403.9
Non-labor Income	3796	3223.504	13095.09	0	390000
Years of Education	3796	13.23	2.33	0	17
Married	3796	0.768	0.422	-	-
Nonwhite	3796	0.307	0.461	-	-
Age	3796	39.76	8.45	25	61

Notes: (i) Data is from the 1996 PSID and includes male heads of household between the ages of 25 and 61. Respondents who report being retired, fulltime students, disabled, full time housekeepers, or in prison were dropped. Respondents were also dropped if they worked fewer than 100 hours in 1996, reported no earnings, or had wages of less than $2 per hour. *(ii)* Hours of work is defined as weeks worked on respondent's main job times hours of work per week. *(iii)* The wage variable is total labor earnings of respondent in 1995 divided by his annual hours of work. *(iv)* The non-labor income measure includes rent, interest, dividends, retirement payments, help from relatives, and lump sum payments of the head as well as wife's income from unemployment, dividends, trust funds, other assets, child support, and miscellaneous non-labor income.

D. Empirical Specification

The labor supply specification is a life-cycle consistent variation of the familiar semi-log specification:

$$\text{Hours}_i = a_1 \ln(\text{Wage}_i) + a_2(\text{Nonlabor Income}_i) + a_3(\text{age}_i) + a_4(\text{age}_i)^2 + \varepsilon_i. \quad (4.3)$$

This specification has several advantages. It implies that wage and income elasticities vary systematically in the sample with hours worked, wages, and non-labor income, allowing for unique elasticities to be calculated for each individual. This specification can be estimated from a single cross-section, and it can also be interpreted in light of life-cycle considerations. Lastly, its familiar form allows comparisons with many previous studies. I estimate equation 4.3 by 2SLS, instrumenting for the wage and for non-labor income.

In order to calculate the income elasticity of labor supply, the income variable needs to provide variation in income that is independent of changes in the wage rate. Although transfer payments such as welfare and unemployment compensation are non-wage income, they are directly contingent on the amount of market work performed in the period of interest. My measure of non-labor income therefore includes income of the head of household from rent, interests, dividends, retirement payments, help from relatives or other sources, and other miscellaneous lump sum payments. The measure will also include the wife's income (where applicable) from unemployment, dividends, interest, trust funds, other assets, child support, and miscellaneous non-labor income. The sensitivity of estimates with respect to other definitions of non-labor income will be considered in the next section.

The income elasticity is given by

$$\frac{a_2 * (\text{Nonlabor Income}_i)}{\text{Hours}_i}. \tag{4.4}$$

Using the Slutsky equation, the compensated wage elasticity is calculated

$$\frac{a_1 - a_2 * \text{Wage}_i * \text{Hours}_i}{\text{Hours}_i}. \tag{4.5}$$

Substituting the measure of non-labor income, wages, and these elasticities into equation 4.1 and assuming $u_{cl} = 0$ gives the coefficient of relative risk aversion:

$$\frac{a_2(\text{Nonlabor Income}_i + \text{Labor Income}_i)}{a_2 * (\text{Labor Income}_i) - a_1}. \tag{4.6}$$

E. Results

The first stage regression results are found in Table 27. Table 28 presents the 2SLS estimation of equation 4.3. Wages are instrumented with education, education^2, and age$*$education and non-labor income is instrumented with a dummy variable indicating whether the head is married. It is important that the instruments be strong; IV estimators perform poorly in finite samples with weak instruments. One test of weak identification is the Cragg-Donald (1993) F statistic. Using the critical values given in Stock and Yogo (2005), the null hypothesis of bias of more than 20% is rejected at the 5% level. In addition, the F-test of the excluded instruments in both first stage regressions is passed easily.

For instrument validity, the model must also be identified. The Anderson (1984) canonical correlations test is a test of whether the excluded instruments are correlated with the endogenous regressors. The likelihood ratio statistic is 25.94 and the null hypothesis of underidentification is rejected with a p-value of less than 0.0001. The Anderson-Rubin (1949) test confirms that log wages and non-labor income are jointly

Table 27. First Stage Regressions of a Life-Cycle Consistent Labor Supply Function: Log Wage and Non-labor Income. (Standard Errors)

	Log wage	Non-labor Income
Variable	Coefficient	Coefficient
Head Age	0.032**	-555.027*
	(0.01)	(252.11)
Years of Education	-0.098**	-3919.247***
	(0.04)	(874.15)
Age^2	-0.000***	1.240
	(0.00)	(2.54)
Education^2	0.005***	106.905***
	(0.00)	(27.07)
Education*Age	0.002***	46.625***
	(0.00)	(9.98)
Married	0.208***	-24.826
	(0.02)	(500.52)
Constant	1.578***	31259.561***
	(0.37)	(8716.01)
Observations	3796	3796
F-statistic	193.033	28.899
Partial R^2	0.19	0.03

Notes: *(i)* * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. *(ii)*Data is from the 1996 PSID and includes male heads of household between the ages of 25 and 61. Respondents who report being retired, fulltime students, disabled, full time housekeepers, or in prison were dropped. Respondents were also dropped if they worked fewer than 100 hours in 1996, reported no earnings, or had wages of less than $2 per hour. *(iii)* Hours of work is defined as weeks worked on respondent's main job times hours of work per week. *(iv)* The wage variable is total labor earnings of respondent in 1995 divided by his annual hours of work. *(v)* The non-labor income measure includes rent, interest, dividends, retirement payments, help from relatives, and lump sum payments of the head as well as wife's income from unemployment, dividends, trust funds, other assets, child support, and miscellaneous non-labor income.

Table 28. Two Stage Least Squares Estimates of Life-Cycle Consistent Labor Supply Function: Total Hours Worked in 1995. (Standard Errors)

Variable	**Coefficient**	**(Std. Err.)**
Log Wage	764.949***	(129.422)
Non-labor income	-0.055***	(0.015)
Age	-32.711*	(17.983)
Age^2	0.347	(0.219)
Intercept	1079.62***	(374.133)
N		3796
$F_{(4,3791)}$		11.742
Sargan statistic		4.243
Anderson canonical correlations statistic		25.940

Notes: (i) * Indicates significance at the at the 10% level, ** at the 5%, *** at the 1% level. *(ii)* Wage is instrumented with education, education^2, and age$*$education, and non-labor income with a dummy variable indicating whether the head is married. *(iii)* Hours of work is defined as weeks worked on respondent's main job times hours of work per week. *(iv)* The wage variable is total labor earnings of respondent in 1995 divided by his annual hours of work. *(v)* The non-labor income measure includes rent, interest, dividends, retirement payments, help from relatives, and lump sum payments of the head as well as wife's income from unemployment, dividends, trust funds, other assets, child support, and miscellaneous non-labor income.

significant in equation 4.3 with a p-value of less than 0.0001. Valid instruments must also be uncorrelated with the error term and correctly excluded from equation 4.3; the Sargan statistic tests for both as a joint null hypothesis. A rejection of the null is cause for doubt as to instrument validity. Table 28 shows a Sargan statistic of 4.243. I fail to reject the null hypothesis of instrument validity with a p-value of 0.12.

In the labor supply regression the coefficient on log wage is positive and significant at the 1% level. The mean compensated wage elasticity is 1.38. The coefficient on non-labor income is also significant at the 1% level, but its magnitude is small. For every additional $1000 of non-labor income, 55 fewer hours per year are worked on average. The mean income elasticity is -0.091.

Chetty uses average income elasticity, wage elasticity, labor income, and non-labor income to compute γ. Based on a number of prominent labor supply studies, Chetty finds γ to be within the range of [0.15, 2.3]. Applying the same methodology by plugging the estimated values of a_1, a_2, and the means in my sample into equation 4.6, I find the coefficient of relative risk aversion for the representative agent to be 0.801.

Chetty provides estimates of γ for a representative agent. Since my estimation is done at the micro level, there is a unique wage elasticity, income elasticity, and risk aversion coefficient for each individual in the sample. I obtain risk aversion measures for the 104 individuals who did not report their educational attainment by using the income that they do report and the coefficient estimates given in Table 28. Summary statistics for the risk aversion coefficients and elasticities are presented in Table 29. The estimates of γ range from 0.048 to 5.992 with a mean value of 0.73.

The instruments and the definition of non-labor income in this study are open to question. As such, it is important that the estimates be fairly robust to changes in these two factors. When I perform a Hausman test by including the predicted

Table 29. Summary Statistics for Labor Supply Elasticities and Chetty's Measure of Relative Risk Aversion

Variable	Obs	Mean	Std. Dev.	Min	Max
Income Elasticity	3900	-0.091	0.376	-9.905	0
Compensated Wage Elasticity	3900	1.382	0.997	0.309	22.491
Relative Risk Aversion	3900	0.733	0.267	0.048	5.992

Notes: I obtain risk aversion measures for the 104 individuals who did not report their educational attainment by using the income that they do report and the coefficient estimates given in Table 28. This brings the total number of observations up to 3900.

residuals from both first stage regressions in the labor supply equation, I am able to reject the null hypothesis that ordinary least squares provides consistent estimates with a p-value of less than 0.001. While the Hausman test indicates that instruments are necessary, the estimates of γ do not change significantly even if I do not instrument for wages and non-labor income.

To check the robustness of the estimates to the instruments used, I re-estimate equation 4.3 using different instruments for non-labor income and the wage. Estimates of elasticities and γ from these robustness checks can be found in Table 30. Different ethnic groups may have different levels of knowledge about investment opportunities which may lead to differences in non-labor income. In 1984, blacks held only .08 percent of bonds and money market funds and only .03 percent of stocks and mutual fund shares (Brimmer 1988), so an indicator variable for whether the respondent is non-white will be included in Columns 1,4, and 5. I include state unemployment rate in Column 3 and 4 because higher rates of unemployment may drive down wages in a particular state. In Column 5 I check whether including higher order polynomials in age and education affect the estimates of γ. The elasticity and risk aversion measures are quite robust to changes in the instruments used. The mean risk aversion estimates

Table 30. Estimated Coefficients on Log Wage and Non-labor Income and Corresponding Means of the Income Elasticity, Compensated Wage Elasticity, and Relative Risk Aversion Resulting from 5 Alternative Sets of Instruments for Log Wage and Non-labor Income.

Estimate	Instruments				
	(1)	(2)	(3)	(4)	(5)
	Married, Nonwhite	Married	State Unemployment Rate, Married	State Unemployment, Married, Non-white	Educ^3,Age^3, age^2*educ, educ^2*age, Non-white
Wage Coef.	771.447	764.949	759.343	764.453	749.413
Income Coef.	-0.054	-0.055	-0.057	-0.056	-0.052
Mean $\varepsilon_{l,y}$	-0.089	-0.091	-0.094	-0.093	-0.085
Mean $\varepsilon_{l^c,w}$	1.372	1.382	1.416	1.409	1.315
Mean RRA	0.728	0.733	0.742	0.739	0.724

Notes: Instruments shared by all 5 specifications are education, education^2, and education$*$age. Column 2 is the specification underlying the estimates reported in Table 28

stay within the range of 0.72 to 0.74 as I change the instruments. Including head's father and mother's education and an indicator for whether the head grew up in poverty also has little effect on the estimates of γ.

There is less consensus on the empirical definition of non-labor income than on the choice of instruments. To test the sensitivity of estimates to changes in the definition of non-labor income, in Table 31 I consider the measure used in Pencavel (2002), namely rent, interest and dividend income of the head. I also use a modified Pencavel measure including rent, interest, and dividend income of the head and wife (where present). Using these measures gives a smaller income elasticity estimate and a larger wage elasticity estimate, but the estimates of γ change little (the mean increases from about .73 to .79). I also consider a measure of non-labor income that follows the same definition underlying Table 27 but also includes the labor income of the wife.

Table 31. Estimated Coefficients on Log Wage and Non-labor Income and Corresponding Means of the Income Elasticity, Compensated Wage Elasticity, and Relative Risk Aversion Resulting from 4 Alternative Definitions of Non-labor Income.

Estimate	Non-labor Income Measure			
	Definition Underlying Table 2	Rent, Interest, &Dividends of Head	Rent, Interest, &Dividends of Head & Wife	Income Measure from Table 2 plus Wife's Labor Income
Wage Coef.	764.949	727.193	727.30	1156.189
Income Coef.	-0.055	-0.090	-0.087	-0.038
$\varepsilon_{l,y}$ Mean	-0.091	-0.048	-0.049	-0.332
$\varepsilon_{l^c,w}$ Mean	1.382	2.005	1.933	1.251
RRA Mean	0.733	0.792	0.785	0.762

Notes: The non-labor income definition underlying Table 2 includes rent, interest, dividends, retirement payments, help from relatives, and lump sum payments of the head as well as wife's income from unemployment, dividends, trust funds, other assets, child support, and miscellaneous non-labor income.

Table 32. Estimated Coefficients on Log Wage and Non-labor Income and Corresponding Means of the Income Elasticity, Compensated Wage Elasticity, and Relative Risk Aversion Resulting from Changing the Restrictions Imposed on the Age and Wage Variables.

	Restrictions				
	(1)	(2)	(3)	(4)	(5)
Age Restriction	None	$25 \leq Age \leq 61$	$30 \leq Age \leq 61$	$20 \leq Age \leq 61$	$25 \leq Age \leq 61$
Wage Restriction	None	Dropped if $wage < 2$	$Wage = 2$ if $wage < 2$	$Wage = 2$ if $wage < 2$	$Wage = 2$ if $wage < 2$
Wage coef.	705.55	764.949	691.682	804.93	776.931
Income coef.	-0.038	-0.055	-0.047	-0.059	-0.057
$\varepsilon_{l,y}$ mean	-0.076	-0.091	-0.086	-0.095	-0.094
$\varepsilon_{l^c,w}$ mean	1.03	1.382	1.231	1.434	1.409
RRA mean	0.654	0.733	0.728	0.723	0.731
Observations	4416	3900	3445	4268	3965

Notes: (i) Column 2 is the specification underlying the estimates in Table 3. *(ii)* Unrestricted data is from the 1996 PSID and includes male heads of household. Respondents who report being retired, fulltime students, disabled, full time housekeepers, or in prison were dropped in all specifications. Respondents were also dropped if they worked fewer than 100 hours in 1996 or reported no earnings.

As discussed in the data section, the sample is limited to individuals who are between the ages of 25 and 61. Respondents who reported having wages of less than \$2 per hour were also dropped. The coefficient estimates are not particularly sensitive to the age constraints, as shown in Table 32. I also examine the sensitivity of the estimates to the wage constraints, by eliminating them or replacing the wage value with \$2 if wages were less than \$2. With neither an age constraint nor a wage constraint, the mean estimate of γ is 0.654. As I change the value of the constraints, γ ranges from 0.723 to 0.733.

F. Further Examination of the Risk Measure

1. Comparing the Chetty and PSID Measures of Risk

Chetty's measure of the curvature of utility over wealth using the PSID data is robust to changes in instruments and alternative definitions of non-labor income. But how does it match up with measures of risk constructed from other domains? The

PSID's risk supplement affords me the opportunity to compare γ with an alternative formulation for 3900 individuals. Since the measure in the PSID asks respondents to consider a job as their only source of income, implying that non-labor income would be equal to zero, the analysis in this section will be done by setting non-labor income equal to zero in the Chetty specification. None of the results change significantly if I allow non-labor income to be positive.

The Barsky et al. (1997) experimental method yields a mean RRA coefficient of 4.584, while the mean of Chetty's measure, γ, with non-labor income equal to zero is 0.673. If I assume that utility is additively separable in consumption and leisure when in reality it is not, γ may be understated, which could explain some of the difference in magnitude of the two estimates. But if the expected utility model is correct and these two methods are measuring the same phenomenon, the rate at which an individual's marginal utility diminishes and his answers to the hypothetical gambles questions should place him at roughly the same point in the distribution of individuals by risk preferences.

To test this, I constructed a variable that was equal to one if the individual's estimated risk aversion was greater than the median according to Chetty's measure, and another if the individual was more risk averse than the median according to the PSID measure which is based on the respondent's answers to the hypothetical gambling question. If these two methods of measuring risk aversion quantify the same phenomenon, we would expect a large overlap between the people in the high category for both measures. Table 33 shows that this is not the case. In fact, the correlation between the two "more risk averse" groups is 0.0375. People that the PSID measure labels "more risk averse" are spread equally between the more and less risk averse categories according to the Chetty measure.

γ looks at how individuals' hours of work respond to changes in their wage or

Table 33. Chetty Measure versus PSID Numerical Measure

	Chetty's Measure		
PSID Measure	$RRA < Median$	$RRA > Median$	$Total$
$RRA < Median$	937	864	1801
$RRA > Median$	1013	1086	2099
$Total$	1950	1950	3900

Notes: The variable "$RRA < Median$" is equal to one if the individual's estimated risk aversion was greater than the median according to a given measure. If the two methods of measuring RRA quantify the same phenomenon, we would expect a large overlap between the high categories for both measures.

their non-labor income. This should tell us something about their preferences. To allow individuals in different risk aversion groups as classified by the PSID measure to respond differently to the same levels of wages and non-labor income, equation 4.3 was estimated separately for each of the 6 PSID risk categories and γ was calculated for each individual using the coefficient estimates for their group. Non-labor income was set to zero in calculating γ, but wages and annual hours worked were allowed to vary by individual. If the PSID measure and γ capture the same aspects of risk aversion, the most risk averse PSID category should have higher values of γ than any other category. The group that accepted all income gambles should have the lowest values of γ. Table 34 shows how estimates of γ vary by PSID risk category.

Individuals who rejected all income gambles and thus are in the most risk averse category according to the PSID measure have an average γ of 0.23, significantly smaller than that of any other group. If the PSID measure and γ captured the same risk tolerance for each individual, γ should decrease monotonically as we move from the most risk averse category to the least risk averse category. Instead, it *increases* until the second least risk averse category. Allowing for different responses to wages

Table 34. Estimates of γ by PSID Risk Category: Wages and Hours Worked Allowed to Vary by Individual

γ for PSID group						
Risk Category	Obs.	Mean.	Std. Dev	Std. Err	Min	Max
Most risk averse	924	0.232	0.111	0.004	0.009	0.877
2^{nd} most risk averse	626	0.611	0.131	0.005	0.090	0.967
3^{rd} most risk averse	544	0.704	0.227	0.010	0.188	3.418
3^{rd} least risk averse	487	0.754	0.120	0.005	0.061	0.957
2^{nd} least risk averse	441	0.750	0.132	0.006	0.031	0.983
Least risk averse	209	0.518	0.157	0.011	0.048	0.931

Notes: Equation 4.3 was estimated separately for each of the 6 PSID risk categories and γ was calculated for each individual using the coefficient estimates for their group. Non-labor income was set equal to zero. If the two measures explain the same phenomenon, the largest values of γ should be in the most risk averse category.

and non-labor income by PSID risk category does not change the fact that the two measures give very different values of risk aversion.

Wages and non-labor income vary systematically with responses to the PSID risk question; individuals who rejected all income gambles and are in the most risk averse category have the lowest average wages and non-labor income. Individuals who accepted all gambles and are in the least risk averse category have relatively low wages, but the highest values of non-labor income on average. We would like for the measure of risk aversion to be capturing differences in preferences and not simply capturing the differences in income between the six groups.

Table 35 examines this issue by estimating equation 4.3 for each of the six risk categories and obtaining the wage coefficient, non-labor income coefficient, compensated wage elasticity, income elasticity, and risk aversion measure, γ, for each group.

Table 35. Estimates of γ by PSID Risk Category: Wages and Hours Worked Set Equal to the Overall Mean

	PSID Risk Category					
Estimate	Most risk averse	2^{nd} most risk averse	3^{rd} most risk averse	3^{rd} least risk averse	2^{nd} least risk averse	Least risk averse
Wage Coefficient	372.48	571.02	367.16	728.38	906.72	770.55
Non-labor income Coef.	-0.003	-0.028	-0.016	-0.076	-0.094	-0.024
$\varepsilon_{l^c,w}$	0.213	0.749	0.458	1.691	2.097	0.779
$\varepsilon_{l,y}$	-0.004	-0.039	-0.023	-0.107	-0.133	-0.034
γ	0.220	0.660	0.642	0.808	0.807	0.558

Notes: Equation 4.3 was estimated separately for each of the 6 PSID risk categories and γ was calculated for each group using the coefficient estimates for the group. Wage income and annual hours worked were set equal to the overall mean in estimating the elasticities and γ. Non-labor income was set to the group mean for estimating Income elasticity and set equal to zero in estimating γ. If the two measures explain the same phenomenon, the largest values of γ should be in the most risk averse category.

Wage income and annual hours worked were set equal to the mean for the entire sample in estimating the elasticities and γ. Non-labor income was set to the overall mean for estimating income elasticity and set equal to zero in estimating γ. The only differences in the measures of γ for the groups are driven by differences in the group's reaction to changes in wages and non-labor income, not differences in wage and non-labor income levels. Even when the only difference between the estimates of γ for the 6 groups are the wage and non-labor income coefficients, the most risk averse PSID category still has the smallest γ. As seen in Table 34, γ is largest for the second and third least risk averse categories.

Under the assumption of constant relative risk aversion, the categorical responses to the PSID risk question place bounds on the measure of relative risk aversion. It is then possible to see if the estimated value of γ for each individual falls within the risk aversion bounds implied by his responses to the risk question. Only 11.2% of the sample did so. Ninety-nine percent of the 11.2% agreement was attributable to the second most risk loving group whose implied bounds were between 0.306 and 1, the range in which 98% of the estimates of γ fall. It seems that the measure of the curvature of utility over wealth implies very different values of RRA than does the

PSID measure. These results bolster Chetty's claim that his paper "provides new evidence that the conventional expected utility model falls short of explaining choices under uncertainty in many domains" (2006, p. 2).

2. Comparisons of RRA by Health Insurance Status

As discussed in the previous section, the PSID measure and the Chetty measure imply very different coefficients of relative risk aversion for the same individual. In order to shed some light on this issue, I examine an individual's health insurance status. In his seminal paper, Pratt (1964) states that an individual has a larger coefficient of relative risk aversion if and only if "he would be willing to pay more for insurance in any situation." I will compare the estimates of RRA for those with health insurance to those without under the two measures to see which of the methods finds higher risk aversion in individuals who choose to insure.

Consumers can choose their level of health insurance coverage in a variety of ways. They can choose jobs which offer health insurance or purchase coverage through a private health insurance company. According to the U.S. Census Bureau, 60.4% of the population of the United States in 2003 had health insurance through employer provided programs. People who are very risk averse likely take health insurance coverage into account when choosing their career. They may be more likely to take jobs with generous health insurance packages and less likely to go into careers where coverage is not offered.

Of the population not covered by employer provided insurance, 15.6% were without health insurance in 2003 and 8.2% chose to purchase health insurance on their own. The remaining 15.8% of the population was covered by government health insurance programs. Basic theory suggests that individuals with health insurance will be more risk averse than individuals who do not have coverage.

Table 36. Comparisons of Means of 1996 Variables: Sub-sample for Which Health Insurance Is Available versus Original Sample

Variable	Original Sample	Sub-sample for whom Insurance Information is available	Sub-sample for whom Insurance Information is not available
	Mean	Mean	Mean
Hours worked	2222.536	2251.240	2160.077
Wage	18.014	18.747	16.419
Non-Labor Income	3196.445	3601.073	2316.017
Years of Education	13.232	13.475	12.695
Married	0.765	0.798	0.693
Non-White	0.303	0.261	0.394
Age	39.684	39.821	39.388
Number of Observations	3900	2672	1228

Notes: (i) The original sample is from the 1996 PSID and includes male heads of household between the ages of 25 and 61. Respondents who report being retired, fulltime students, disabled, full time housekeepers, or in prison were dropped. Respondents were also dropped if they worked fewer than 100 hours in 1996, reported no earnings, or had wages of less than $2 per hour. *(ii)* The sub-sample is made up of the individuals from the original sample who reported their health insurance status in 2001. *(iii)* Hours of work is defined as weeks worked on respondent's main job times hours of work per week. *(iv)* The wage variable is total labor earnings of respondent in 1995 divided by his annual hours of work.

Information on health insurance coverage is not provided in the 1996 PSID. Respondents were, however, asked to provide this information in 2001. Because of the panel nature of the data, I can link the respondents in 1996 to their health insurance information in 2001[3]. Unfortunately, in 1997 the PSID moved to biennial data collection and the number of families in the core sample was reduced by a third. Of my original 3900 observations, I was able to match 2672 individuals with their 2001 information. Table 36 compares the means of individuals from the 1996 total sample to the sub-sample for which health insurance information was available. The remaining sample looks similar to the original sample. The individuals for whom I have health insurance information have slightly higher wages, more non-labor income, and are more likely to be white.

I do not consider those with government provided health care as part of the analysis (98 observations). This leaves me with a sample of 2574 individuals. Of

[3]Re-estimating RRA using 2001 data is not an option as the PSID risk measure is only included in 1996.

Table 37. Summary Statistics by Health Insurance Coverage Group

	No health coverage	Has health coverage
Variable	Mean	Mean
Hours worked	2051.238	2273.153
Wage	10.889	19.609
Non-Labor Income	1365.032	3820.666
Yrs Education	11.991	13.636
Married	0.573	0.82
Non-White	0.414	0.243
Age	38.345	39.836
Number of Observations	220	2354

Notes: (i) The sub-sample is made up of the individuals from the original sample (as described in Table 1) who reported their health insurance status in 2001. *(ii)* Ninety-eight individuals with government provided health insurance were dropped. *(iii)* Hours of work is defined as weeks worked on respondent's main job times hours of work per week. *(iv)* The wage variable is total labor earnings of respondent in 1995 divided by his annual hours of work.

this group, 86.56% has employer provided health coverage, 4.90% has private health insurance purchased directly, and 8.55% is not covered by a health care plan. Sample statistics for the covered and uncovered group are presented in Table 37. Individuals with health coverage work more, earn a higher wage, have more non-labor income, and more education. They are a year and a half older on average, more likely to be married, and more likely to be white.

Individuals with no health insurance have a mean value of γ of 0.584 as compared to a mean of 0.761 for those with health coverage. The p-value of the test of equal means is less than 0.0001. According to the PSID measure, the mean RRA for those without health insurance is 4.066 and for those with insurance is 4.583. We can reject that these two groups have the same RRA; the p-value on the test for equal means is 0.003. The results from both measures accord with our expectation that more risk averse individuals are more likely to insure.

As discussed in the previous section, we do not want differences in γ to be simply capturing the differences in income between the two groups. To shed some light on this issue, I run separate regressions for individuals with health insurance and those without. This allows the two groups to respond differently to the same levels of wages and non-labor income, which should capture any differences in preferences between the two groups. I then evaluate the coefficient of relative risk aversion at the mean for the entire sample of individuals in 1996 so that the estimates will not be driven by the fact that those with health insurance generally have higher wages and more non-labor income. Evaluating γ at \$42,054.40 of labor income and \$3,610.78 of non-labor income, I find γ for those with no health insurance to be 0.923 and 0.785 for those with health insurance. Bootstrapping this estimate, I find a mean difference of -0.133 with a standard error of 3.05. When I employ this implementation of the Chetty method, I find no significant difference in the relative risk aversion of individuals with health insurance as compared to those without. Due to the small sample of individuals without health insurance, the precision of the estimates was poor which could explain the lack of difference of the two estimates.

G. Conclusion

The explanations for many interesting economic phenomena employ the concept of risk aversion, with the Arrow-Pratt measure as the generally accepted standard. Raj Chetty (2006) introduced a new formula for estimating RRA from labor supply information, which was applied to micro data in this paper. I compare this measure to the measure included in the 1996 PSID which is based on hypothetical gambles over lifetime income. The two measures appear to be substantially different, which is compatible with Chetty's claim that the expected utility model cannot fully explain choices under uncertainty. It seems that expected utility explains little, if any, variation in risk preferences. I further investigate the differences in these two measures by utilizing the fact that individuals who choose to insure should be more risk averse than those who do not. I compared estimates of RRA for those with health insurance and those without, ceteris paribus. In the PSID sample, the hypothetical gambles method supported this hypothesis, with higher levels of risk aversion in those with health insurance coverage. The Chetty method also supported this hypothesis when income was allowed to vary for each individual. When I allow individuals with different health insurance status to respond differently to wages and non-labor income and set income levels at the sample mean, the Chetty method finds no significant difference between the two groups.

CHAPTER V

CONCLUSION

Women are found in systematically different careers than are men. It is difficult to determine whether this is because women have the primary responsibility for the home and children and thus desire different career attributes than men do or because women face discrimination in the labor market. The lack of widely available data on a woman's expectations about her future make investigating this problem particularly challenging as expectations are vitally important to understanding this phenomenon. If a woman expects to have a large family or many in-home responsibilities, she may be more likely to choose a career that is compatible with this choice. Expectations are difficult to measure and are not always fulfilled and ignoring their importance can lead one to incorrect conclusions about the cause of a woman's career path.

Different religious groups have different beliefs on the importance of child bearing and the division of labor within the home. The Southern Baptist church, for example, teaches that "a wife is to submit herself graciously to the servant leadership of her husband" and she "has the God-given responsibility to respect her husband and to serve as his helper in managing the household and nurturing the next generation." Because of differences in religious ideology, fertility expectations should differ by religious denomination. If this is the case, including a woman's religion in regressions on job attributes should control for fertility expectations and expected effort expended in household chores. In general, it appears that Jehovah's Witnesses are likely to choose the most flexible careers followed by Pentecostal, Catholic, and Baptist women. There is no significant difference in the flexibility of Protestant groups which are more likely to stress gender equality and women with no religious preference. When Jewish women differ from women with no religious preference, they tend to be in less flexible careers. These results are consistent with the human capital notion that women are choosing different careers than men rather than being forced into different job paths.

Women appear to be choosing jobs that allow them to take responsibility for home production and child rearing. These choices of wives are likely to have an effect on their husbands as well. Male wage regressions that include marital status dummy variables find marriage wage premiums of between 10 and 40%. This could be precisely because wives are taking the primary responsibility for home production and husbands are free to focus their attention on productivity at work. On the other hand, perhaps factors unobserved to the researcher may both make a man more productive and more likely to marry. It is also possible that employers discriminate against single men. To investigate this issue, I again use religious denomination, this time as a proxy for specialization within the home. Pentecostal men and women, for example, were not as likely to agree that "when a husband and wife both work, housework should be shared equally" as were Catholic husbands and wives (Brinkerhoff and MacKie 1984). When a Pentecostal man marries, his wages increase dramatically, even after controlling for a number of observable characteristics. Catholic men, Protestant men, and Baptist men also benefit from large wage increases after marriage. The wage premium for Jewish men appears to be in the form of increased wage growth. There is no reason to expect that employers would prefer married Pentecostal men to single Pentecostal men more than they prefer married Catholic men to single Catholic men, so discrimination cannot explain these findings. Fixed effects estimation removes time-invariant individual characteristics, and it is difficult to imagine a scenario in which selection into marriage differs across religious denomination by a characteristic that varies over time. These significant differences in marriage premium by religious denomination provide new support for the productivity hypothesis.

The choice of a career, a level of specialization within the home, whether to marry and when, and most other important life decisions are heavily dependent on one's tolerance for risk. The role of risk preferences in these life choices is not fully understood, largely because risk aversion is very hard to empirically quantify. Raj Chetty (2006) introduced a new formula for estimating RRA from labor supply information, which I apply to micro data from the 1996 Panel Survey of Income Dynamics. The 1996 PSID also contains a measure of risk aversion derived from individual's responses to hypothetical gambles over lifetime income. I compare the two measures and find them to be substantially different, which is compatible with Chetty's claim that the expected utility model cannot fully explain choices under uncertainty. I further investigate the differences in these two measures by utilizing the fact that individuals who choose to insure should be more risk averse than those who do not. I compared estimates of RRA for those with health insurance and those without, ceteris paribus. In the PSID sample, the hypothetical gambles method supported this hypothesis with higher levels of risk aversion in those with health insurance coverage. The Chetty method also supported this hypothesis when income was allowed to vary for each individual. When I allow individuals with different health insurance status to respond differently to wages and non-labor income and set income levels at the sample mean, the Chetty method finds no significant difference between the two groups.

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