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Empirical Essays on Health Care for Children and Families

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Empirical Essays on Health Care for Children and Families

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Dedicated to my family

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Empirical Essays on Health Care for Children and Families

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This dissertation consists of three empirical essays investigating different aspects of health care for children and families. The first essay examines the effectiveness of adherence to American Academy of Pediatrics guidelines for preventive pediatric health care. Using a national longitudinal sample of children age two years and younger, we investigate whether compliance with prescribed periodic well-child care visits has beneficial effects on child health. We find that increased compliance improves child health. In particular, higher compliance lowers future risks of fair or poor health, of some history of a serious illness and of having a health limitation.

The second essay examines child health care utilization in relation to maternal labor supply. We test the hypothesis that working-mothers trade off the advantages of greater income against the disadvantages of less time for other valuable tasks, such as seeking health care for their children. This tradeoff may result in positive, negative, or no net impacts on child health investment. We estimate health care demand regressions that

include separate variables for mother's labor supply and her labor income. Our results indicate that higher maternal work hours reduce child health care visits; higher maternal earnings increase them. In addition, wage-employment, as opposed to self-employment, is detrimental to child health investment. A further finding is that preventive care demand for younger children is less sensitive to maternal time and income changes. We also find that detrimental time effects dominate beneficial income effects.

The third essay studies intra-household resource allocation as it pertains to its demand for preventive medical care. We test the income-pooling hypothesis of the common preference model by using individual specific medical care consumption data and present evidence on the allocation of household resources to the medical needs of the child, husband and wife. Our results are in line with the findings of previous studies that emphasize the ongoing importance of the traditional gender role of woman as the primary caregiver. We find that the resources of the wife have a greater positive impact on child's and her own preventive care demand than does the resources of the husband. In contrast to most studies from developing countries, we find that US families do not exhibit differential health care demand based on child gender. It is also noteworthy that the wife's education level has a greater positive impact than that of her husband does on both the husband's and her own preventive care utilization.

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Chapter 1

Introduction

This dissertation consists of three separate chapters, each of which is briefly summarized below. All three chapters empirically investigate different aspects of health care for children and families.

Chapter 2 examines the effectiveness of adherence to prescribed periodic well-child care visits. It is important to identify the good investment interventions in child health. In this study, we use a longitudinal national sample of children two years of age and younger to assess the productivity of routine well-child visits. The American Academy of Pediatrics authoritatively structures the well-child care in the U.S., producing guidelines that specify the frequency and the content of them. Despite a general recognition of the benefits of compliance with the recommended visits, there exists a lack of evaluation of the efficacy of fulfilling these recommendations. Accounting for potential endogeneity of preventive care utilization, we investigate whether compliance with well-child care visits has beneficial effects on child health. We find that increased compliance improves child health. Controlling for initial health and common risk factors, higher compliance lowers future risks of fair or poor health, of some history of a serious illness and of having a health limitation.

Chapter 3 investigates child health care utilization in relation to maternal labor supply. The mother's employment has potential implications for virtually all aspects of child well-being; investment in child health is no exception. Since mothers are the primary arrangers of health services for their children and usually accompany them to health care providers, child health care usage may depend significantly on whether and how much their mothers work outside the home. This study investigates whether increasing scarcity of the mother's time affects her child's health investment, measured by preventive and curative care utilization. We test the hypothesis that working-mothers must trade off the advantages of greater income against the disadvantages of less time for other valuable tasks, such as seeking health care for their children. This tradeoff may result in positive, negative, or no net impacts on child health investment. We estimate health care demand regressions that include separate variables for mother's labor supply and her labor income. The findings suggest that higher maternal hours of work correlate with reductions in child health investment, while increases in mother's earnings are associated with higher investment in child health. We also find that wage-employment, as opposed to self-employment, is detrimental to child health investment. A further finding is that the preventive care demand for younger children is less sensitive to maternal time and income changes. Net total impact of maternal work is found to be negative.

Chapter 4 examines how resource-controlling power within the household affects the intra-household distribution of preventive medical care, while testing the unitary model of household behavior. This study, besides providing substantiated evidence on this issue for a developed country, is different from most previous intra-household resource allocation studies in its use of individual specific medical care consumption data, which allows estimation of separate demand functions for children, wives and husbands. Our results are in line with the findings of previous studies that emphasize the ongoing importance of the traditional gender role of women as the primary caregiver. We find that the mother's non-labor income has a greater positive impact on child preventive care demand than does non-labor income of the father. We do not find evidence of either gender bias or parental gender preference in child health investment. Furthermore, we find that the wife's non-labor income and her education level have a greater positive impact than those of her husband do on her own preventive care utilization. The wife's education has a greater positive impact on the husband's preventive care demand than does his own education level. Non-labor incomes of spouses are not significant determinants of husband's preventive care demand.

Chapter 2

Child Health: Does Compliance with Well-Child Care Visits Make a Difference?

2.1 INTRODUCTION

The importance of preventive childhood care has long been emphasized at the national level, through such programs as the Maternal and Child Health Services Block Grant, the State Children's Health Insurance Program and the Medicaid's Early and Periodic Screening, Diagnostic and Treatment Program. Well-child care (WCC) represents the main source of preventive health care for children in the United States. It consists of regularly scheduled visits aimed at disease prevention and health promotion through immunizations and through health education; early detection and treatment of disease through physical examinations and via disease-specific screening; and the provision of anticipatory guidance in all aspects of child rearing.

Traditionally, the American Academy of Pediatrics (AAP) authoritatively guides the structure of WCC in the U.S. Since 1967, the AAP has produced *Recommendations for Preventive Pediatric Health Care*, also known as the periodicity schedule, based on observed practices and expert opinion. This document details a timeline for scheduled well-child visits from birth through 21 years of age and specifies the preventive services that should be provided at each well-child visit. Since its introduction, the AAP offered modifications several times, most of which broaden the obligations of primary care physicians. The number of recommended WCC visits also has expanded rapidly over time, growing from 14 in 1967 to 28 recommended visits today. The current AAP recommendations call for 6 well-child visits by 12 months of age (at 1, 2, 4, 6, 9, and 12 months), another 3 visits evenly distributed through the second year of life, 1 annual visit from 3 to 6, 1 visit every other year from ages 7 to 10 and 1 visit per year from 11 to the age of 21.

Many national public programs widely incorporate the AAP *Recommendations for Preventive Pediatric Health Care*. Bright Futures, a major initiative launched by The Maternal and Child Health Bureau to improve health promotion and preventive services for children, encourages compliance with the AAP guidelines. Many state Medicaid programs' early periodic screening, diagnosis, and treatment services frequently draw directly from the recommendations of the AAP in setting their schedules for periodic screening. In addition, many private health insurers base their WCC benefits on the AAP periodicity schedule.

A substantial body of research has investigated the WCC utilization rates and factors associated with obtaining recommended WCC (Bryd et al.,1999; Freed et al.,1999; Ronsaville et al.,2000; Yu et al.,2002; Selden, 2006). Focus mainly centers on finding ways in which to increase the adherence to WCC visits. The rate of compliance with WCC visits remains a policy target for health promoters, which indicates the value society places on the AAP guidelines.

Despite a general recognition of the benefits of compliance with WCC visits, there exists a lack of evaluation of the efficacy of these recommendations' fulfillment. With the exception of immunizations, the clinical effectiveness of adherence to prescribed periodic WCC visits has not been demonstrated (Gadomski et al.,1998). The purpose of this study is to investigate the impact of compliance with well-child visits on the health of children of age two years and younger. We focus on this age group, because the highest concentration of preventive services usage appears in this group. Using nationally representative longitudinal data from the Medical Expenditure Panel Survey (MEPS), we investigate whether or not increased compliance has beneficial effects on child health.

There are studies in the Pediatric field that reveal correlations between WCC visit compliance and better health outcomes. Hakim and Bye (2001) found that, among Medicaid-enrolled children, compliance with the AAP recommended series of WCC visits during the first two years of life related to fewer avoidable hospitalizations. Hakim and Ronsaville (2002) demonstrated a positive effect of compliance in lowering the risk of emergency department use. One clinical study (Gilbert et.al.,1984) administered a randomized trial in 570 healthy newborns. It compared infants receiving an average of 7.6 visits in the first two years of life with infants averaging 4.8 visits in the first two years of life, finding no significant differences in their health outcomes. This study, however, did not examine explicitly the effects of receiving the AAP recommended series of WCC visits.

Assessing causality between compliance and health outcomes is not trivial. In principle, a randomized controlled trial that eliminates all forms of spurious causality could establish of WCC productivity. However, conducting these experiments remains expensive and they are subject to ethical concerns. Moreover, since they are conducted on a population limited by geography, by race or by some other aspects, results may not generalize. In observational studies, on the other hand, demand for WCC potentially is endogenous to health status. One needs to distinguish between the effects of compliance on health and the effects of health on compliance. Current research evidence indicates that children in poor health are more likely to comply with WCC visits (Selden, 2006). This selection effect may make a substantial contribution to the association between compliance and health. To investigate the effects of compliance on health, while minimizing the contribution of selection effects, we analyze longitudinal observational data and evaluate the relationships between initial compliance and subsequent change in health, while controlling for initial health. This, however, may not totally account for the possibility of reverse causality. We employ an Instrumental Variable (IV) approach to further control for the potential endogeneity. The findings indicate that increased compliance with WCC visits improve health trends of children. In particular, when controlling initial health and common risk factors, a higher compliance level leads to lower future risks of fair or poor health, of having some history of a serious illness and of having a health limitation.

In the next section, we present a conceptual framework to guide the empirical specification. Section 3 describes the empirical strategy and the results are presented in Section 4. Section 5 concludes the study.

2.2 CONCEPTUAL FRAMEWORK

The analysis is based on estimating a health production function that describes the relationship between combinations of health inputs and the resulting health outcome. Consider a simple two period model. In the first period, parents decide whether to take their child to regularly scheduled WCC visits, as recommended in the periodicity schedule. Through timely intervention and/or anticipatory guidance, such routine visits may prevent future health problems. The second period reveals a child's health status and the parents attain utility from their child's health status. If the WCC visits are productive, children in greater compliance with the periodicity schedule should be less likely to suffer a health shock.

Formally, we assume the child is endowed with a stock of health H_1, in the first period, which depreciates to H_2 in the second period. H_2 is defined as

$$H_2 = f(M_1)H_1$$

M_1 is a variable indicating the degree of compliance with WCC visits. The function $f(M_1)$ characterizes the depreciation of health from period 1 to period 2. We assume that

$$\ln f(M_1) = \alpha + \beta M_1 + \delta X + \mu + \varepsilon$$

where X is a vector of individual level, household level and community level observable characteristics affecting health depreciation. Child's health endowment μ, represents effects of unmeasured factors known to the parental decision maker (i.e. child-specific unobserved heterogeneity in health). The change in health stock from period 1 to period 2 is given by

$$\ln H_2 - \ln H_1 = \alpha + \beta M_1 + \delta X + \mu + \varepsilon \qquad (1)$$

Household utility in each period depends on the child's health status and consumption level in each period,

$$U(H_t, C_t) \quad t = 1,2$$

The household's budget constraint is given by

$$p_m M_t + p_c C_t = I_t$$

where C_t represents consumption of a composite good. I_t is exogenous income; p_m and p_c are money prices of preventive care visits and consumption, respectively.

Since this is a two period model, $M_2 = 0$ and $C_2 = I_2$. Maximization of the utility function subject to the health production function and resource constraint yields the following reduced form demand function for M_1:

$$M_1^* = M(p_m, p_c, I_1, X, \mu) \tag{2}$$

For a given level of initial health H_1, observables X and health endowment μ, substituting M_1* into the equation (1) yields the health production function

$$H_2^* = H(H_1, M^*{}_1, X, \mu) \tag{3}$$

In equation (2), the conditioning of preventive care demand on the child-specific unobserved health heterogeneity implies that estimates of the health production function that do not account for the endogeneity of parental preventive care use decisions are biased and inconsistent.

2.3 Empirical Strategy

We estimate a health production function that links health inputs to child health outcome. For each child in the sample, two consecutive years of data are available. Assuming that benefits of preventive care are not immediate, we characterize the production function of the child's health in the current period as a function of his/her health status in the previous period, augmented by health inputs in the previous period. The estimated health production function is of the following form:

$$H_2 = a + bH_1 + cM_1 + dX + e$$

where H_2 and H_1 are the child's health outcome measure in period 2 and period 1 respectively; M_1 measures the amount of WCC received by child; X is a vector of observable characteristics that affect child's health outcome in period 2 and e is the error term that also includes child's unobserved characteristics.

The previous discussion suggested that endogeneity complicates the estimation of causal effect of WCC adherence. Not only does the use of preventive services affect health outcome but also the anticipated outcome may affect the demand for preventive care. Often, the inclusion of initial health status as an explanatory variable is used as a method of controlling for the possibility of reverse causation or health selection. It is possible, however, that, there will still be a health selection problem. Such selectivity bias can be avoided by implementing an IV approach in which child health outcome is estimated jointly with a behavioral model for which health inputs themselves are choices. To isolate the effect of preventive care that is independent of unobserved child characteristics, the IV approach requires one or more instrumental variables, representing observable factors that influence preventive care use while not directly affecting child health outcomes.

2.3.1 Data

Data for this study come from the Medical Expenditure Panel Survey (MEPS) sponsored by the Agency for Healthcare Research and Quality, and the National Center for Health Statistics. The MEPS, which began in 1996, is an annual, nationally representative survey of households. It consists of three interrelated surveys: the Household Component, the Medical Provider Component and the Insurance Component. For this study, we use data from the MEPS Household Component. The set of households selected for the Household Component is a sub-sample of those participating in the National Health Interview Survey. At each interview, the questionnaire accrues information about each household member, and, one designated household respondent reports all of the MEPS data. The survey consists of an overlapping panel design in which a new panel is initiated each year and is followed for two years via five in person interviews.

Certain populations of particular policy interest are over-sampled in the MEPS to allow adequate sample sizes for examination. Furthermore, the sample design of the survey includes stratification, clustering and multiple stages of selection. The complex survey design of the MEPS needs to be taken into account by applying MEPS survey weights to produce estimates and by using an appropriate technique to derive standard errors associated with the weighted estimates.

A series of calendar year-specific MEPS data files are produced annually. These files contain data from the second year of a continuing panel in conjunction with the first year of a new panel. The person identifier remains the same for a participant for his/her entire duration in the survey. The survey provides detailed information on demographic characteristics, health status, utilization of medical services, health insurance coverage, income and employment for every member of the household. In terms of medical care services usage, the survey collects data about all hospital (emergency room, inpatient and outpatient) events, physician services, home health care and prescribed medicines. The MEPS asks respondents to keep a calendar of medical visits they make during the year and to supply supporting paperwork regarding those visits. Visit-level files that include detailed characteristics associated with that visit are produced each year.

For purposes of this study, we pool two consecutive years of data on children of age two years and younger from panel 6, 7 and 8. The MEPS assigns family identification numbers to individuals in the survey; while we construct family units and assign family characteristics to each child, we exclude children whose mothers are not present in the household. This yields a sample of 2,958 children. However, as some children do not have complete information on some health outcome measures, we use a subset of this full sample that has complete information for these measures.

In this study, we examine the second year impact on child health outcome of first year WCC compliance. Table 2.1 summarizes the weighted descriptive statistics of the variables used in the analysis.

2.3.2 Dependent Variables

Measures of child health. Using the MEPS household-reported health status data, we construct three measures of child health in the second year: (1) a binary variable for whether the child is in fair or poor health; (2) a binary variable for whether the child suffered any serious illness at any time prior to the survey interview; (3) a binary variable for whether the child has any physical, sensory (e.g. vision, hearing), cognitive limitations.

Parents are asked to rate their child's health on a five-point scale: excellent, very good, good, fair or poor health. The first health status measure is a binary variable coded to 1 for children whose health was described as "fair" or "poor," as opposed to "excellent," "very good," or "good." Self-rated health status is an increasingly common measure of health in empirical research (Ettner, 1996; Saunders, 1996; Deaton and Paxson, 1998; Smith, 1999). This is supported by literature that shows that it predicts mortality and morbidity (Okun et al., 1984; Connelly et al., 1989; McCallum et al., 1994; Idler et. al., 1995). In their review of several studies, Idler and Benyamini (1997) found that self-rated health was an independent predictor of mortality while controlling for a broad array of health status indicators and correlates; they concluded that self-ratings represent an "irreplaceable dimension of health status". Self-assessed health also has been found to predict functional decline and chronic disease (Idler and Kasl, 1995). Furthermore, it has been shown that a continuous health status measure constructed from a categorical self-reported health variable highly correlates with other continuous measures of health (Gerdtham et al.,1999). Lundberg et al. (1996) also demonstrated good overall reliability of self-assessed health.

The MEPS asks respondents how true or false is the statement: "Child *has never been* seriously ill". Parents then rate this statement as being definitely true, mostly true, do not know, mostly false, or definitely false. The second binary variable takes the value 1 if the parent reported this statement to be "mostly false" or "definitely false," indicating that the child *has been* seriously ill at some point before the survey interview. The survey also asks whether the child in question has any activity, functional, cognitive or sensory limitations. The third health measure indicates the presence of such limitations. Adequate

WCC care may prevent such limitations by providing early intervention with the detection of problems causing such limitations.

In the second year of the survey, 1.9 percent of children were in fair or poor health, 22.1 percent had a serious illness prior to the second-year survey and 2.4 percent were limited in some way.

2.3.3 Explanatory Variables

Compliance ratio. The key explanatory variable of interest is the compliance ratio that measures adherence to WCC visits in the first year of the survey. It is the number of completed WCC visits during the first year as a percentage of age-specific recommendation from the AAP.

Well-child care is identified using visit-level information both on office and hospital outpatient visits. The MEPS asks the primary reason for the visit and categorize each visit based on the type of medical service received. For this analysis, visits are coded as WCC if the primary reason given was “general check-up”, “well child exam” or “immunizations or shots”.

Criteria for WCC visit compliance are based on the 2000 AAP periodicity schedule for 9 designated visits during the first two years of life. As previously mentioned, this schedule prescribes 6 visits by 12 months of age (at ages 1, 2, 4, 6, 9, and 12 months) and another 3 visits by 24 months of age (at 15, 18 and 24 months). To construct the compliance ratio, we merged the AAP periodicity schedule into the MEPS using the child’s age (in months) at the completion of the survey’s first year. The compliance ratio is expressed as a percentage throughout the analysis.

On average, children completed 66 percent of the recommended WCC visits in the first year. Figure 2.1 shows the histogram of the compliance rate in the first year; the histogram peaks at the values equal to 0 and 100.

Number of WCC visits. The compliance measure incorporates both the number of and the timeliness of the visits. In order to assess whether it is the adherence to a scheduled series of visits that yields health benefits, rather than the mere frequency of visits, we re-estimate all health outcome models by replacing compliance ratio with the number of WCC visits in the first year.

Initial health. To account for the reverse causality between compliance and health, we control for the initial health status at the time of the first-year survey. Regression equations for each dependent variable defined above include the first lag of the dependent variable as an explanatory variable. Controlling for the child's earlier health status results in a modeling of the change in health over a two-year period, adjusting for the baseline level of health. The fact that the health indicators in this study are self-reported provides yet another reason to control for initial health. There may be significant differences in the way in which people, whose true health actually is the same, respond to the survey questions. Nevertheless, as long as these differences remain constant in both periods, they will not affect the estimates.

In each year the MEPS asks households if they have any medical condition and collects detailed information on reported conditions. To further control for the baseline child health, each model also includes the number of reported health conditions in the year 1.

In the first-year survey, 2.2 percent of the children were reported to be in fair or poor health, 20.3 percent had a serious illness at some point and 4.1 percent had a health limitation. The average number of health conditions reported in the year 1 is 2.7.

Other explanatory variables. Other control variables include characteristics of the mother: age in years, education in years and employment status. Also included in the models are controls for child's sex, age in months, family income, family size, an indicator for whether the father is present in the household, race/ethnicity indicators (Hispanic, black and white) and an indicator for metropolitan statistical area (MSA) residence. All explanatory factors are measured at the first-year survey.

The average age of children in the sample was 19.2 months. Slightly more than half of the children were male, 21.6 percent were Hispanic and 14.5 percent were black. Average annual family income was $48,640 and average family size was approximately 4. Sixty-four percent of children had working-mothers. Mothers' average education was 12.8 years and their average age was 29.3 years. Approximately, 30 percent of children did not have their father present in the household. Eighty-two percent of children lived in an MSA.

2.3.4 Instrumental variables

By controlling for initial health status, we seek to avoid the reverse causality problem between compliance and health. It is possible, however, that an endogeneity problem will still exist. Some uncontrolled component of health status or some unobserved factor that affects both the compliance and the rate of depreciation of health over time may lead to further endogeneity issues. To test for this, we adopt an IV approach, which estimates child health outcomes jointly with a behavioral model in which the compliance level is itself a choice variable.

In medical care effectiveness studies, variables that describe cost of and access to care are used widely as instruments for medical care utilization. Following the literature, we use indicators for insurance status (any private insurance during the year, only public insurance during the year and uninsured for the entire year) and census region (Northeast, Midwest, West and South) as instruments for the amount of WCC received.

Insurance status reflects both monetary costs of and the adequate access to preventive care. Previous research enumerated marked differences among census regions and divisions in well-child visit compliance (Olson et al., 2005; Selden, 2006). These differences remain significant even after controlling for a wide array of child and family characteristics. One explanation could be geographic differences in the supply of pediatricians providing preventive care. Another could be geographic variation in the monetary cost of medical care. There is evidence that compliance may be related to the prices paid to providers for well-child care (McInerny, 2005).

2.3.5 Estimating the Effectiveness of Compliance with Well-Child Care Visits

To assess the effects of compliance on child health while minimizing confounding due to health selection, we analyze the relationships between initial compliance level and subsequent change in health, while, controlling for initial health. Specifically, the dependent variable in each regression equation is a health measure assessed in year 2 and the independent variables are the health measure assessed in year 1, the initial compliance level along with other factors known to affect child health. We then specify regression equations as linear probability models; not only because it allows the most flexible way to handle instrumental variables, but also because the coefficient estimates are easy to interpret. We have confirmed, however, that the marginal effects from probit specifications evaluated at the mean yield nearly identical results to the linear probability model.

Controlling for initial health status may not totally account for the possibility of reverse causality between compliance and health. To account for this, we adopt an IV approach. For each health outcome equation, specification tests are performed, testing for the validity of employed instruments and for the endogeneity of the compliance measure. If the compliance measure establishes itself as endogenous, the Two-Stage Least Squares (2SLS) method is used to obtain unbiased and consistent estimates.

This study adjusts all its estimates for the sample design of the MEPS including stratification, clustering, multiple stages of selection and disproportionate sampling, using survey weights. Standard errors for weighted estimates are adjusted for complex survey design. The survey weights also allow accounting for the non-independence of individual observations stemming from participation by more than one child per household.

2.4 EMPIRICAL RESULTS

2.4.1 The Risk of Being in Fair or Poor Health

Tables 2.2 through 2.5 present the estimation results. Table 2.2 presents the results predicting fair or poor self-rated health in the year 2. The findings suggest that compliance with well-child visits is important for child health. The compliance ratio is highly significant with a negative sign, implying that higher compliance in the first year reduces the risk of fair or poor health in the second year, conditional on initial health. A one-percentage point increase in the compliance ratio reduces the probability of reporting fair or poor health in the year 2 by 0.0001. This may appear to be a very small impact. However, when compared to the overall sample risk of being in fair/poor health, it turns out to have a considerable effect. For example, the results indicate that the risk of being in fair/poor health for a child who is up-to-date with WCC visits (100 percent compliance) is lower by 0.005 than the same risk for a child who made only half of the recommended visits (50 percent compliance). Given that the overall risk of fair/poor heath in the sample is about 2 percent, this represents a 25 percent reduction in that risk.

In order to determine if IV estimation is necessary, we test for whether the compliance measure correlates with the unobservable factors. Indicators for insurance status and census region are chosen as instruments. The bottom panel of Table 2.2 presents specification test results. We regress the compliance ratio on the instruments and the other regressors defined in the previous section. An F test of joint significance of the instruments in this first stage regression provides a check for instrument relevance. Rejection of the null shows relevancy of the instruments. The Hansen J Statistic for the overidentification test for all instruments reflects instrument exogeneity. A rejection of the null casts doubt on the validity of instruments.

Test results presented in Table 2.2 confirms the validity (relevance and exogeneity) of employed instruments: (1) We reject the null hypothesis that the instruments have no joint effect in the first stage regression (F statistic= 16.45, p value= 0.00); and (2) We are not able to reject the null hypothesis that the instruments were jointly valid; the instruments which were excluded from the second-stage regression passed the test of overidentifying restrictions (test statistic= 2.17, p value= 0.82).

We test for endogeneity by using the Rivers-Vuong approach. In the first stage, the compliance ratio is regressed on the instruments and the other explanatory variables. In the second stage, the health outcome variable is regressed on the residual from the first stage and on the other explanatory variables. A significant coefficient on the residual implies that compliance measure is endogenous. The first stage regression results are presented in the Appendix A.

The predicted residual from the first-stage regression was not significant when included in the health outcome regression; the null hypothesis of exogeneity of compliance measure cannot be rejected (test statistic= 0.18, p value= 0.67). This implies that IV estimation is not warranted; the Ordinary Least Squares (OLS) estimation provides an appropriate estimation method in this case.

Regarding the other explanatory variables, initial health status variables are highly significant and with the expected signs. Children in fair or poor health and those with more health conditions in the year 1 are more likely to be in fair or poor health in the year 2. The results also suggest that Hispanic children exhibit significantly worse health trends[1]. The other explanatory variables, however, do not have a significant impact on the change in health over this two-year period. For this age group, biological factors may play a larger role in health determination than socioeconomic factors.

2.4.2 The Risk of Having a History of a Serious Illness

The estimation results for the probability of a serious health event history prior to the second-year survey are shown in Table 2.3. The first column presents OLS estimation results. Conditional on initial health status, higher compliance reduces the risk of a history of a serious health event. This effect is marginally significant at the 10 percent level. Specification test results for this health outcome, presented in the bottom panel of Table 2.3, confirm that employed instruments are valid; that is, they both are relevant (F

[1] Objective indicators of poor health such as self-reported illnesses primarily explain lower health ratings by Hispanics. When we include indicators for whether the child has a history of a serious illness and whether the child has any health limitation as explanatory variables, the coefficient for Hispanic indicator turns to be insignificant.

test statistic= 12.60, p value=0.00) and are exogenous (F test statistic=8.24, p value= 0.14). The Rivers-Vuong test statistic rejects the null hypothesis of exogeneity of compliance ratio (test statistic=6.23, p value= 0.03), implying that IV estimation is more appropriate than OLS estimation for this health outcome variable.

The second column of Table 2.3 depicts the IV estimation results for the probability of a history of a serious health event prior to the second-year survey, conditional on initial health. The estimated coefficient of the compliance ratio is negative and highly significant after accounting for endogeneity. The estimated coefficient is much larger than the OLS coefficient. Specifically, a one-percentage point increase in the compliance ratio decreases the probability of reporting some history of a serious illness by 0.0027, conditional on initial health. This is a considerable decrease when compared to the overall risk in the sample. To illustrate, the risk of reporting some history of a serious illness for a child who is up-to-date with WCC visits (100 percent compliance) is lower by 0.135 than the same risk for a child who made only half of the recommended visits (50 percent compliance). The sample mean risk is about 0.20. This corresponds to a decrease in sample mean risk of about 68 percent.

Initial health variables have significantly positive coefficients as expected. Higher levels of mother's education significantly reduce the future risk of a serious illness, conditional on initial health. Children with absent fathers have worse health trends. The other variables are not significant.

2.4.3 The Risk of Having a Health Limitation

Estimation results for the probability of reporting the presence of some health limitation in the second year are presented in Table 2.4. The compliance ratio has a negative coefficient and is significant at the 5 percent level. Conditional on initial health, a one-percentage point increase in the compliance ratio decreases the future risk of a health limitation by 0.0002. For example, the risk of having a health limitation for a child who is up-to-date with WCC visits (100 percent compliance) is lower by 0.01 than the same risk for a child who made only half of the recommended visits (50 percent compliance). Given that the mean risk of having a health limitation in the sample is about 3 percent, this represents a 33 percent reduction in that risk.

Specification test results presented in the bottom panel of Table 2.4 confirm that the chosen instruments are relevant (test statistic= 13.77, p value=0.00) and exogenous (test statistic=6.32, p value= 0.20). The Rivers-Vuong test statistic, however, does not reject the null hypothesis of exogeneity of compliance (test statistic=0.37, p value= 0.56). We therefore rely on the OLS estimation results for this health outcome.

Regarding the other control variables in the model, variables for initial health status have significantly positive coefficients. The risk of any health limitation declines with age in months, conditional on initial health. The other explanatory variables do not have a significant impact on change in health.

2.4.4 The Number of Well-Child Care Visits

Table 2.5 shows the results obtained by replacing the compliance ratio by *the number of WCC visits* in the first year. For each of the three health outcomes, specification tests (results are not shown for brevity) could not reject the exogeneity of the number of WCC visits; hence OLS estimates are reported in Table 2.5. More WCC visits reduce the risks of adverse health outcomes in each of the three models; however, these effects hold no significance in most cases. A marginally significant (10 percent) effect is found only for the probability of having a health limitation. This suggests that it is the act of compliance with a scheduled series of preventive visits that yields health benefits, rather than the mere frequency of preventive visits.

2.5 Summary and Conclusion

Government agencies, professional organizations and health insurance-related organizations stress the importance of compliance with the prevailing guidelines for the WCC visits. They have hypothesized that adherence to periodic WCC visits reduces childhood morbidity, disability and mortality while promoting optimum growth and development. Despite a general recognition of the benefits of adherence to periodic WCC visits, the clinical effectiveness of adherence has not been demonstrated.

This study analyzed longitudinal panel data for a nationally representative sample of children two years old and younger to assess the effectiveness of compliance with the AAP recommended WCC visits. To avoid reverse causality between compliance and child health, we controlled for the initial health status of children and employed an instrumental variable approach when there were further endogeneity issues present. The findings suggest that compliance with the current AAP periodicity schedule improves health trends of children. In particular, we find that, controlling for initial health and common risk factors, higher compliance lowers future risks of being in fair or poor health, of having a history of a serious illness and of having a health limitation. The estimated effects imply considerable reductions in the risk of adverse health outcomes when compared to the overall mean risks in the sample. Another implication is that the beneficial effects of higher compliance are timely: this year's higher compliance yields beneficial effects in the next year. One could conjecture that the health benefits of compliance continue beyond the next year and accumulate over time. Unfortunately, the data do not allow us to address this possibility. The analysis also suggests that the compliance with a scheduled series of preventive care encounters is the key to the positive outcome, not the mere frequency of preventive visits.

The estimated beneficial effects of compliance were robust to several sensitivity checks. The use of different health status measures provides the first sensitivity check; in each of the three models of health outcomes, higher compliance had beneficial effects on child health. Other sensitivity tests monitored whether the results were sensitive to the measure of compliance. Well-child visits may occur slightly before or after the recommended date with little or no reduction in their effectiveness. We allowed compliance ratios to exceed 100 percent in order to minimize the impact of such small variations in the timing of visits. Results are robust to this specification. Defining broader time ranges for construction of the compliance ratio, to allow for variability in visit timing, also did not change the results.

This study provides further support of the role of early, periodic health supervision in preventing pediatric morbidity. National efforts to improve child health should focus on increasing compliance with periodic preventive care. To this end, policy makers, researchers and clinicians should work toward eliminating systemic barriers to comprehensive and timely preventive care for children associated with the greatest disparities in service delivery and quality. They should also support efforts to educate families on the importance of the periodic WCC visits, overcome language and other cultural barriers, provide for after-hours care and remind parents of upcoming appointments. It is encouraging to note that this research, which demonstrated the productivity of compliance over a two-year period, suggests that policies to ensure adequate compliance are likely to provide an immediate impact in improving child health.

Given the societal costs of unhealthy children, benefits of WCC are worthy of further exploration. The estimated effects of compliance in this study included the joint influences of the number and the timing of WCC visits in preventing adverse health outcomes. These, however, are only proxies for the concept of complete preventive care and do not provide a measure of the clinical content and its quality. It is important to explore the mechanisms by which WCC visits might prevent adverse outcomes, identifying components of the WCC visits that are especially effective. Another worthwhile extension to this study would be investigating different health outcomes using different data. Finally, the effectiveness of WCC for older children is also an important issue that deserves analysis.

Chapter 3

Maternal Employment and the Demand for Children's Health Care

3.1 INTRODUCTION

The increase in labor force participation among married women, especially women with children, marks one of the most significant changes in the United States during the 20th century, an evolution leading to important changes in familial organization. Employment among all women rose from 43 percent to 60 percent between 1970 and 2001 (U.S. Bureau of Labor Statistics). As of 2002, 78 percent of married women with school-age children and 63 percent of married women with infants and preschoolers seek employment outside the home (U.S. Census Bureau). Employment rates are even higher for single mothers. Because mothers play a central role in the care and nurturing of children, the substantial growth in mothers' labor force participation naturally led to questions about how households with working-mothers cope with increased maternal time constraints—in particular, changes they may make in the ways children are raised and cared for and the subsequent effects on children's well-being.

Several previous studies investigated maternal employment effects on a wide variety of child well-being indicators, including health and nutritional status, cognitive abilities, behavioral problems, school achievement and emotional adjustment. Despite numerous studies on the topic, there still exist many unanswered questions about the impact of maternal employment on child well-being. Among them is whether it has any effect on child health investment, in particular preventive and curative health care services rendered to children. Since the mother is the primary decision maker as to where and when such services will be obtained and usually also accompanies the child to the provider of health services, changes in female market and non-market roles may affect child health care utilization.

Maternal work may affect the health care utilization of children through a number of channels. In this study, we focus on the increasing lack of mothers' time and the extent to which this time deficit affects parents' health care demand for their children. Women are the principle arrangers of health services for their children and their spouses. Family health responsibilities traditionally are seen as an integral part of women's role as the mother and wife. As women, particularly mothers of preschool and school age children, increasingly enter the labor force the time available to them for traditional family activities diminishes. Their joint roles as caregivers and as providers of family income may conflict with one another. Increased hours allocated to market work may crowd out other valuable uses of mothers' time, such as seeking medical care for children. On the other hand, additional income they bring into the household may help to ensure increased investment in child health. In this study, we hypothesize that a tradeoff exists between detrimental time allocation effects of mother's work and beneficial effects coming through changes in family income. The net effect of maternal labor supply is then an empirical question.

Among the various factors that affect pediatric care demand, the effects of maternal labor market behavior have not received much attention. To our knowledge, no studies to date investigate child's health care utilization in relation to maternal labor supply. Most studies examining the health care demand for children focus on other factors while controlling for subsets of maternal employment (Colle et al., 1978; Vistnes et al., 1995). It is important to understand maternal time allocation effects on child health investment because, with the presence of extended insurance coverage, time constraints might become an increasingly important determining factor. Expansion of health insurance should lessen the effect of monetary costs of pediatric care and increase the relative importance of time costs.

Investigation of maternal labor supply effects on child health investment may also shed light on recent findings in child health research. A growing body of literature has found deleterious effects of maternal employment on child health (Anderson et al. 2003; Berger et al. 2005; Crepsinek and Burstein 2004). Mediating factors, however, are not fully known. One potential mechanism may be the decrease in child health investment due to increasing scarcity of mother's time. The time spent by women in seeking preventive and curative care is an important input into the production of child health. Mothers who work may lack the time to adequately perform this role.

Using a nationally representative sample of children, this study investigates the impact of maternal labor supply on the likelihood and the level of preventive and curative care utilization among children. The findings suggest that higher maternal work hours reduce child health care visits, while higher maternal earnings increase them. The analysis also reveals that wage-employment, as opposed to self-employment, is detrimental to child health investment. The net total impact of maternal work is found to be negative, indicating that the negative effects of time dominate the positive income effects.

The data used in this study do not allow accounting for the effect of unobserved factors that are potentially correlated with both maternal labor supply and with child health investment. We employ two approaches to address this problem. First, we use observed data on maternal and child characteristics to proxy unobserved factors to the fullest extent; second, we use an informal test to assess the importance of the problem of unobserved heterogeneity.

The paper's organization is as follows: In Section 2, we present a conceptual framework that will guide the empirical investigation. Section 3 describes the empirical specification and the data. Results are presented in Section 4 and Section 5 concludes.

3.2 CONCEPTUAL FRAMEWORK

The conceptual framework for the empirical investigation assumes child dependence on parents or other adults to seek and obtain health care and that the characteristics of the children, the parents, and the family all influence the health care children receive. As suggested by the literature on the demand for health care, we expect child health status, insurance coverage, age, sex, race and ethnicity, family income, family size, parental education and location of residence to play a role in child health care utilization. In addition, parents' demand for their children's health care is expected to be influenced by maternal employment. Mothers' employment yields potential implications for virtually all aspects of child well-being, and health investment is no exception. As mothers play a dominant role as gatekeepers to the health care system, the demand for children's health care may depend significantly on whether and how much their mothers work outside the home. Time spent by mothers in seeking preventive and curative medical care for their children is an important input into the production of child health. Time spent in the market work erodes time available for the mother to perform her gatekeeper role in pediatric care, in arranging health services for her children and in escorting them to health care providers. Assuming that no caretaker would be more motivated than the child's mother to arrange health care; other things equal, children of working-mothers might get less use of health care than children of nonworking-mothers.

On the other hand, maternal labor market activities contribute to household resources that may increase parents' health care demand for their children. As a result, a positive income effect may offset a potentially negative time allocation effect. Furthermore, if preferences of mothers and fathers differ, such that mothers are more inclined to spend their income in ways beneficial to the children, the effects from mother's earned income will be greater than from other household income sources. Several studies have found that consumption of goods that enhance child well-being, such as education and health, rise with the share of household income under maternal control (Thomas, 1990, 1994; Phipps and Burton, 1998).

This framework suggests that households in which mothers supply market work must trade off the advantages of greater income against the disadvantages of less time to attend to children's health care needs. This tradeoff may result in positive, negative, or no net impact of maternal work on child health investments. For this study, no a priori hypotheses are formed with regard to mother's employment and child health care utilization. We expect the loss of time available for attending child health care needs and the gain in income to work in opposite directions, with the net effect to be determined empirically.

This research leads to several hypotheses regarding the effects of time allocation and income. First, we expect that the nature of the maternal work itself to be an important factor in consideration of time allocation effects of maternal employment. Self-employment usually is considered more compatible with childcare than wage-employment, providing greater flexibility in hours or the possibility of combining work and childcare. Self-employment may mitigate potentially detrimental effects on child health investment associated with reductions in the mother's allocation of time to household activities.

Second, we expect time and income effects of maternal work to yield less impact for younger children, especially infants, than for older children. Younger children have strong needs for adequate medical care, especially in their early life developmental stages. Adequate health care services are most essential for them. We hypothesize that the essential nature of the demand for younger children's health care renders it less sensitive to maternal time and income changes.

Last, the data used in this study allow us to analyze the relative impact of maternal earnings on child health investment. In addition to estimating the potentially conflicting time allocation and income effects of maternal work, we also investigate the differences in the impacts of maternal earnings versus other income in the household.

3.3 EMPIRICAL STRATEGY

In order to estimate the hypothesized tradeoff between income and time effects, as well as the overall net impact of maternal work on child health care demand, as Glick suggested (1998), we estimate reduced form health care demand regressions that include separate variables for mother's labor supply and her labor income. The following equation for the child health care demand is estimated:

$$HCU = \alpha + \beta_1 H + \beta_2 I + \beta_3 OI + \beta_4 X + e \tag{1}$$

where HCU is a measure of child's health care utilization, H is the mother's hours of work, I is the mother's labor income, OI is other (non-maternal) household income and X is a vector of explanatory variables including child and household characteristics determining the health care demand for the child. We use maternal hours of market work as an inverse proxy for the time available for the mother to attend to the health care needs of her child. Including other household income, OI, as a separate explanatory variable allows analysis of the relative importance of maternal earnings for child well-being.

In this specification, β_1 is the effect of an additional hour of maternal work, controlling for associated changes in income, while β_2 is the effect of changes in maternal labor income, controlling for hours of work. If there is a tradeoff between the detrimental time allocation effect and the beneficial income effect of maternal employment, as hypothesized, β_1 will be negative and β_2 will be positive. The net impact of an additional hour of work is equal to (β_1+ β_2w), where w is the mother's hourly wage rate. Multiplying this expression by H, mother's hours of work, gives the net total impact of maternal work on children's health care utilization. The net total impact of maternal employment will be negative if the time effect is larger in absolute value than the income effect multiplied by the wage rate, given that $\beta_1<0$ and $\beta_2>0$, as hypothesized.

3.3.1 Data and Methods

The data used in this study come from the Medical Expenditure Panel Survey (MEPS), an annual, nationally representative survey of households sponsored by the Agency for Healthcare Research and Quality and the National Center for Health Statistics. The survey consists of an overlapping panel design in which a new panel is initiated each year and followed for two years via five in-person interviews. This ongoing survey has collected information about each household member at each interview since its inception in 1996. The MEPS is a stratified and clustered random sample with weights that produce nationally representative estimates for medical care utilization and a wide range of other health-related and socioeconomic characteristics.

A series of calendar year-specific MEPS data files are produced annually. These files contain data from the second year of a continuing panel, along with the first year of a new panel. The person identifier remains the same for a person for his/her entire duration in the survey. The survey provides detailed information on demographic characteristics, health status, utilization of medical services, health insurance coverage, income and employment for every member of the household. In terms of medical care services usage, the survey collects data about all hospital (emergency room, inpatient and outpatient) events, physician services, home health care and prescribed medicines. The MEPS asks respondents to maintain a calendar of medical visits they make during the year and to supply supporting paperwork regarding those visits. Each year, they produce visit-level files that include detailed characteristics associated with that visit.

For purposes of this analysis, we pool the MEPS data for the years 1997 through 2000. The sample includes children age 12 and younger. The MEPS assigns family identification numbers to each individual in the survey. Using these, we construct family units and assign maternal and family characteristics to each child in the sample. The empirical work focuses on the mother's characteristics and the child's use of health care, given that the mother usually accompanies the child to the health care providers. Thus, the sample is restricted to children who lived with their mothers, yielding a sample of 19,390 observations. We weighted all analyses using weights that reflect both the sample design of the MEPS and survey non-response, and we adjusted standard errors for clustering of observations within children and within families. We estimate multivariate regression models using four different measures of utilization as dependent variables: (1) an indicator of whether the child had any office based-preventive visit during the year;

(2) an indicator of whether the child had any office-based curative visit during the year; (3) total number of office-based preventive visits during the year; (3) total number of office-based curative visits during the year.

The MEPS provides detailed information on each medical visit of survey participants during the year. Office-based visits are visits that transpire in office-based settings or clinics, excluding outpatient, inpatient or emergency visits. We identify the type of care using visit-level information on office visits. The MEPS asks the primary reason for the visit and categorizes each visit based on the type of medical service received. Preventive visits are ambulatory visits identified by the respondent mainly as a well-child exam, an immunization or a general check-up not associated with a specific condition. Curative visits are defined as all remaining office visits. We distinguish between preventive and curative (remedial) services because the decision process underlying the demand for care in each case likely differs. In addition, we examine the likelihood of a visit separately from the volume of use (total number of visits). Since family decisions are most likely to affect the decision to enter the health care system, volume of use is affected by physician decisions as well as family decisions; it is useful to analyze these two measures separately.

The key explanatory variables in the study are the maternal work variables: mother's weekly hours of market work and mother's weekly labor income. Maternal labor supply acts a proxy for time constraints mothers face. The MEPS provides a number of employment characteristics of participants, including weekly hours of work and hourly wage rates, collected at each of the three rounds of the annual interview. We average these across survey rounds to derive "average weekly hours of work" and "average hourly wage". In order to obtain weekly labor income, we multiply average weekly hours of work by average hourly wage. For self-employed mothers, we divide their annual personal incomes by 50 (approximate number of weeks worked in a year), to calculate their weekly earnings. For mothers who do not supply any market work, weekly hours and earnings are coded to zero. In order to construct weekly income accruing to other (non-maternal) household members, we sum the annual personal incomes of each family member (except the mother) and divide by 50. All specifications also include a control for mother's work status.

Consistent with the literature on the demand for health care, other explanatory variables in the models include mother's age and education, as well as indicator variables for the child's age, sex, and race/ethnicity. The models also account for measures of child health status, including self-rated (or parent-rated) general health and having any physical, social, or developmental limitation; insurance status[2], family size, indicators for census region, presence of the father in the household, presence of other adults in the household, an indicator variable for residence in a metropolitan statistical area (MSA) and for whether the mother is the survey respondent.

We use the linear probability model (LPM) to estimate the propensity to use health care, since the coefficient estimates are easier to interpret in this case. We have confirmed, however, that the marginal effects from probit specifications yield similar results to the linear probability model. Since the variables for number of visits take on small, nonnegative integer values, we employ a negative binomial regression to model these variables.

Table 3.1 presents means and standard deviations of the variables used in the analysis. Table 3.2 reports descriptive statistics on maternal labor market participation and type of employment, defining participation as having positive average weekly hours

[2] Addressing endogeneity of insurance status is out of scope of this study. Insurance status proxies for the medical care prices the families face. In addition, maternal employment may affect child health care use through provision of insurance coverage. Controlling for insurance status allow us to focus on partial effects of time allocation and income.

of work. The participation rate of mothers of young children is slightly lower than for mothers of older children. The majority of mothers are in wage-employment.

3.4 EMPIRICAL RESULTS

3.4.1 Children's Preventive Care Utilization

Estimation results for child preventive care utilization are shown in Tables 3.3 and 3.4; Tables 3.5 and 3.6 present the results for children's curative care utilization. Table 3.3 shows LPM estimates of equation (1) when the dependent variable is the likelihood of a child having at least one preventive visit during the year. A one-hour increase in mother's labor supply per week decreases the propensity of a preventive visit by 0.0029. A one-dollar increment in mother's weekly earnings increases the propensity of a preventive visit by 0.0001. As hypothesized, maternal employment appears to involve a trade-off between the advantages of higher income against the disadvantages of less time to attend to child health care needs. The coefficient for mother's work status is statistically insignificant when controlling for mother's work hours and earnings.

In order to test the hypothesis that the adverse effects on child health investment associated with reductions in mother's time will be stronger for wage-employed mothers than for self-employed mothers, separate values are entered for maternal hours of work in self-employment and wage-employment. The results are presented in Column 2 of Table 3.3. Controlling for earnings, increases in wage-employment hours significantly reduce the likelihood of a preventive visit, an effect that is significant at the 1 percent level, whereas mother's self-employment hours are statistically insignificant. Increasing mother's time in wage-employment appears to be detrimental for child health investment, while increasing self-employment hours has no effect.

One of our hypotheses is that the effect of maternal work would be less on younger children. In Column 3, we report results obtained by adding interactions of maternal work variables and child age to the basic model of column 1. We find a significantly positive interaction between maternal work hours and the indicator for the child being under age 2. Given the negative coefficient of the main effect of mother's hours of work, this suggests that the impact of mother's time spent working will indeed be less for children under age 2, when compared to the impact on older children. In addition, we find a significantly negative interaction of maternal earnings and the dummy for the child being under age 2. Given the positive coefficient of the main effect of maternal income, this suggests that the effect of mother's earnings will be smaller for children under age 2 compared to the impact of her income on older ones. These findings confirm that younger children's propensity to receive preventive services is less sensitive to maternal time and income changes.

The hypothesis that maternal resources benefit children more than other family members' resources is tested by comparing the magnitudes of the estimated coefficient of mother's weekly income and the coefficient of other household income. As in the case of maternal income, increases in non-maternal household income associate with a higher likelihood of a preventive visit. However, an F-test of equality of these effects shows that the effect of other family income is much smaller than the effect of maternal income (F-statistic=18.99, p-value=0.00). This finding supports the hypothesis that the household does not pool its income and that mothers have a relatively high propensity to spend their incomes in ways that improve child welfare.

Coefficient estimates found for other explanatory variables are presented in Table B.1 in the Appendix. The presence of an extended family with more available caretakers might offset any decrease in utilization of health care services among children of working-mothers. As expected, the presence of other adults in the household significantly increases the likelihood of a preventive visit. Among other covariates, maternal age and education, having health insurance and the presence of the child's father in the household, significantly increase the propensity to use preventive services. Preschool age children are more likely to use preventive services than school age children are. The insignificant coefficient on the male dummy variable indicates the absence of male gender bias in child health investment. Black children and Hispanic children are less likely to have a preventive visit than their white peers are; while children in fair/poor health or good health are more likely to have a preventive visit compared to children in very good/excellent health. If they have any health limitation, they are more likely to use services. Growth in family size also reduces the propensity to make a preventive visit, as increasing the number of children in the household may place greater demands on maternal time for childcare and restrict the mother's ability to seek health care outside the home. Children who live in the Northeast and the Midwest regions of the U.S. are more likely to use preventive services than are the children in the South. Children living in an MSA are more likely to use preventive services. Regional differences in health care utilization have been reported previously and may be related to physician features as well as differences in health care delivery systems (Welch et al., 1993).

Table 3.4 reports the negative binomial regression results for the annual number of preventive care visits; the estimated marginal impacts are reported in Table B.2 in the Appendix. Findings are very similar to those for the likelihood of a preventive visit. Higher maternal hours of work correlate with reductions in the number of preventive visits; in contrast, increases in mother's earnings associate with higher number of preventive visits, as expected. Results in Column 2 show that increasing work hours in wage-employment significantly reduce the number of preventive visits while self-employment hours do not have a significant impact. Findings reported in Column 3 confirm the hypothesis that time and income effects of mother's work matter less for younger children than for older children. Interactions of mother's hours and child age dummies are significantly positive and interactions of mother's earnings and child age dummies are significantly negative.

The finding that increases in maternal income will benefit the children more than increases in other household income also holds for the level of preventive care use. An F-test for equality of these two effects imply that the coefficient of maternal income is significantly larger than the coefficient of other household income (F-statistic=11.78, p-value=0.00). The estimated coefficients of other determinants of the level of preventive care children receive yield the very same implications as those discussed previously. For the sake of brevity, we do not report them.

3.4.2 Children's Curative Care Utilization

We replicate the analysis for child curative care demand. Table 3.5 reports the results for the likelihood of having at least one curative visit and Table 3.6 reports the results for the annual number of curative visits. Results in Column 1 of Table 3.5 show a significantly negative impact of mother's hours of work and a significantly positive impact of mother's earnings on the propensity to make a curative visit. Including self-employment and wage-employment hours in the regression as separate variables, Column 2, yields a significantly negative effect of wage-employment hours and a statistically insignificant effect of self-employment hours. The results in Column 3 show whether these effects vary depending on children's age. The coefficient of each of the interaction terms is insignificant, implying that for illness-related visits, maternal work effects are the same across age groups. The results also suggest that children of working-mothers are more likely to seek curative care than those of nonworking-mothers are, even after controlling for mothers' time constraints and income.

As in the case of maternal income, other household income significantly increases the likelihood of using curative services; however, the impact of maternal income is larger than the effect of other household income. An F test, however, could not reject the equality of the effects of these two incomes (F-statistic=2.69, p-value=0.11). The estimated coefficients for other covariates (not reported) yield the similar implications as those for the preventive care demand.

Table 3.6 presents the negative binomial estimates of the number of curative visits during a year. Increases in maternal hours of work significantly and adversely affect the level of curative care use. The estimated effect of maternal income is positive and significant as expected. Other household income is not a significant determinant of the level of curative care received. The coefficient for the mother's work status is statistically insignificant. Column 2 of Table 3.6 shows that more stringent wage-employment hours adversely affect child health investment; there is no statistically significant effect of self-employment hours on curative visits. Finally, including interactions of maternal work variables and children's age dummies, Column 3 of Table 3.6, yields no significant variation in the effects of maternal hours and earnings that depends on child's age. The coefficient estimates (not reported) for the other determinants of the level of curative care have the same implications as discussed previously.

3.4.3 Net Effects of Maternal Employment

The net effect of maternal employment on child health investment combines the impacts of mother's hours of work and her earnings. Estimated impacts of maternal hours and earnings from the previous section can be used to measure the net impact of maternal employment, which is given by (β_1+β_2w). Table 3.7 shows the estimated net effects evaluated at the mean hourly earnings of working-women (w=$12) and their standard errors.

Using the results from column 3 of Table 3.4, the net effect of an hour of work on the probability of having a preventive visit is -0.0021. The net total effect of maternal work is equal to the net effect of an hour of work multiplied by mother's work hours. We first calculate the net total effect of maternal employment for each child (whose mother is working) using the mother's actual hours of work and her actual hourly income and then take the average of the calculated net total impacts. The mean net total effect of maternal work on the likelihood of a preventive visit for the sample of working-mothers is -0.076. A similar calculation for the number of preventive visits yields a net marginal effect (evaluated at mean wage rate) of -0.0061 and a net total effect of -0.226. This indicates that, controlling for other factors, the probability of having a preventive visit for children of working-mothers is lower by 0.076 than the same probability for children whose mothers do not work. This is a 16 percent decrease from the sample mean probability of having at least one preventive visit. Children of working-mothers also have fewer preventive visits, by 0.226, than do children of nonworking-mothers. This corresponds to a 23 percent decrease from the sample mean.

With regarding to the total impact on curative care visits, similar calculations yield a net marginal impact of maternal work on the likelihood of a curative visit of -

0.0008 and a net marginal impact on the number of curative visits of -0.0085. The mean net total impacts for the likelihood and the number of curative visits are -0.029 and -0.310, respectively. The decrease in the likelihood of a curative visit is 5 percent of the sample mean and the decrease in the number of curative visits is 15 per cent of the sample mean. These estimates indicate that the negative time allocation effects of maternal work dominate positive income effects.

3.4.4 Addressing Unobserved Heterogeneity

We are not able to control for all possible factors potentially correlated with both maternal employment and child health investment. Addressing these remaining sources of confounding is challenging. Although the MEPS provides two-years of data on each child, there is not enough variation in maternal labor supply over this short period that would allow us to use panel data methods to control for unobserved factors. The instrumental variable estimation requires identification of variables correlated with maternal employment but unrelated to child health investment, other than through maternal employment. It is difficult to enumerate such variables. Instead, we use two alternative approaches to address this problem. First, observed data on maternal characteristics are used to proxy unobserved factors to the fullest extent; second, we use an informal test to assess the importance of unobserved heterogeneity. The informal test centers on the idea that if selection based on observed characteristics is not particularly critical, it seems unlikely that selection on unobserved characteristics is important (Altonji, Elder, and Taber, 2005). The test consists of comparing results from parsimonious and more fully specified models of health care utilization. Showing that the estimated effects of maternal work variables does not change substantially after adding another control variable provides evidence that the effect of unobserved factors is negligible because unobserved factors correlated with this additional variable will be partially (or fully) accounted for (Altonji, Elder, and Taber, 2005; Duncan, 2003; Ruhm, 2004). If unobserved variables correlated with the control contributed to a significant omitted variables bias, we might expect the estimate of the effect of maternal work to change after adding the observed control.

For each regression model estimated, we start with a parsimonious model of health care demand and gradually expand this model to include a larger number of child and maternal characteristics[3]. In all models, the inclusion of additional covariates has little impact on the magnitude of the coefficients for maternal work. Almost all of the additional covariates, however, appear to be important predictors of health care demand. This suggests that selection along unobserved characteristics may not be particularly important because: (1) observed maternal and child characteristics are good predictors of health care demand for children; and (2) the inclusion of these characteristics has little impact on the estimate of maternal employment on health care demand. From this, it is reasonable to conclude that the presented estimates of the effect of maternal labor supply on child health investment will not be greatly biased.

[3] We start with a specification that includes variables related to maternal work (mother's hours of work, her earnings, work status and child's insurance status) and gradually include other child, mother and family characteristics that we discussed previously. For the sake of brevity, I do not report the results of these experiments.

3.5 Summary and Conclusions

In this study, we investigated how increased demands on women's time affect child health investment, measured by preventive and curative care utilization. Since women are generally children's primary caregivers, they take responsibility for seeking preventive and curative pediatric health care. Due to the time constraints faced by working-women, there may be conflicts between their roles as caregivers and as providers of family income. These conflicts have potentially important implications for child well-being. We use maternal hours of work as an inverse proxy for the time available for a mother to attend to the health care needs of her child. We test the hypothesis that families in which mothers work outside the home must trade off the advantages of greater income against the disadvantages of less time for other valuable uses of time, such as seeking preventive and curative care for children. This tradeoff may result in positive, negative, or no net impacts on child health investment.

We estimate health care demand regressions that include separate variables for the mother's labor supply and her labor income. As hypothesized, maternal employment appears to involve a trade-off for child health care demand. Higher maternal hours of work associate with reductions in the likelihood of and the quantity of both preventive and curative visits, while mother's earnings increase them. We also find that wage-employment, which in general provides less flexibility than does self-employment, proves more detrimental to child health investment. A further finding is that preventive care demand for younger children is less sensitive to maternal time and income changes. For curative care, there exist no significant differences in the maternal work effects across age groups. The net total impact of maternal work is found to be negative, signifying that the negative time effects dominate the positive income effects.

The results also suggest that child health investment increases more with maternal income compared to other household income. This supports the findings of other studies which show that income in the hands of women is more likely, relative to spouse's or other family members' income, to be spent on items that benefit children (e.g., health care), suggesting that women have relatively stronger preferences for child welfare.

This study provides a first step toward understanding a relatively unexplored topic: maternal employment effects on child health investment. Future research can extend our work by utilizing panel data methods and instrumental variable approaches for endogenous treatment of maternal work.

Chapter 4

Female Control of Household Resources and the Intra-Household Demand for Preventive Care

4.1 INTRODUCTION

This paper studies intra-household resource allocation as it pertains to its demand for preventive medical care. We test the income-pooling hypothesis of the traditional unitary (i.e. common preference) model by using individual specific medical care consumption data and present evidence on the allocation of household resources to the medical needs of the child, husband and wife.

There are two competing approaches to modeling intra-household resource allocation. The first views the family as a single decision making unit. This "unitary" or "common preference" model assumes that household behaves as if it has one set of preferences, represented by a household utility function. This amounts to assuming that a single utility function is maximized as if all the family members have the same preferences or there is a benevolent dictator that determines allocations (Becker, 1965). Then household pools all family resources and subsequently allocates them according to some common rule. The second approach views a household as a collection of individuals with different resources and divergent preferences. Intra-household resource allocations are seen as bargaining outcomes (McElroy and Horney, 1981), or as Pareto-efficient outcomes achieved through some collective decision-making processes (Chiappori, 1988).

The unitary and the collective household models yield different predictions. The unitary model suggests that the party controlling resources is irrelevant in intra-household resource allocation; since all the household members pool their resources, consumption decisions depend only on aggregated household resources and not on its individual components. Conversely, the common implication of the collective models is that changes in individual resources, by changing household members' decision-making powers, translate into altered household consumption patterns. This holds important implications for public policy, particularly benefit transfer programs. For example, the collective models suggest that welfare benefits have varying effects on children's welfare, depending on which parent receives them. The unitary model, on the other hand, suggests that as long as the amount of transfer itself does not change the named recipient of the benefit is irrelevant to its allocation and impact.

Several studies have attempted to test the validity of the unitary model. The income-pooling hypothesis of the unitary model has been rejected in a variety of country settings in both the developed and developing world. Overall findings in this literature challenge the assumption that couples take an absolute stance on pooling their resources, either by pooling all resources or none at all. For example, results indicate that husbands and wives may pool their income for some goods (e.g. larger purchases including durables) but not for others (e.g. smaller purchases, including non-durables) (Phipps & Burton, 1998).

There seems to be a consensus in the intra-household resource allocation literature that giving control over money to women is a more effective way to increase household well-being, because women have stronger preferences for it. Haddad and Hoddinott (1994, 1995) show that an increase in the wife's control over financial resources is associated with greater expenditures on food, while reducing expenditures on alcohol and tobacco. In addition, several empirical studies that explore the relationship between female decision-making power and child outcomes show that in specific contexts, women are more dedicated to using resources for children's benefit than are men, and that children fare better in households in which the wife has greater control over resources (Thomas; 1990, 2002).

In undertaking our analysis, we test the income-pooling hypothesis of the unitary model. We examine how relative spousal power within the household affects the distribution of a specific good, preventive medical care. In particular, we examine the impact that women have on the preventive care received by family members and shed light on the theoretical implication that emerges from the collective modeling approach, i.e., when women have relatively more resources than those of her spouse, family members utilize more preventive care.

The novelty of this study in the intra-household decision-making literature lies in its use of individual-specific medical care consumption data to examine familial resource allocation. Since individual-specific consumption data distinguishable across household members often proves difficult to find, most of the previous studies looked at the aggregate household consumption of goods. Some household expenditure surveys provide data on total spending on child-specific goods, such as clothing and childcare, but do not provide data on spouse-specific consumption. Consequently, previous studies could examine intra-household resource allocation towards children, but the evidence on the effect of relative control over resources on wives' and husbands' consumption patterns is very rare. In this study, we present empirical evidence on the distribution of household resources toward medical needs of children, of husbands and of wives.

Measuring the relative control over resources always proves to be a complex task. Many studies employ husbands' and wives' incomes as measures of their relative controls; however, income directly relates to labor supply, which itself may be an outcome of their relative controlling power. In order to surmount this problem, previous research focused on economic resources exogenous to labor supply. They include assets, both current and those brought into marriage (Beegle et al., 2001; Thomas et al., 2002; Quisumbing and Maluccio, 2003; Doss, 1999), unearned income (Schultz, 1990; Thomas, 1990) or transfer payments and welfare receipts (Lundberg et al., 1997). Some studies use education attainment as a proxy for marital power; since education highly correlates with potential earnings, an increase in education level can be associated with enhanced control in the determination of household resource allocations.

Following the literature, we use non-labor incomes and completed years of education to measure spouses' relative controlling power and test whether the consumption of preventive care is higher, other things being equal, in families in which wives have greater control over resources. If non-labor incomes and education attainments are good indicators of controlling power, then differences in preferences of men and women should be revealed in differential effects of spousal non-labor income and education on the preventive care usage if individual family members.

It has been argued that tests of income-pooling based on non-labor income may not completely avoid endogeneity issues, since non-labor income may reflect past or current household behavior. Earlier studies used components of non-labor income that are more likely to be exogenous, such as income from physical and financial assets, to mitigate this problem. We take a similar approach and use asset incomes as a measure of spouses' relative control over resources, to check for the robustness of the effects of non-labor incomes.

Our work also shed light on the issue of differential investment in sons and daughters in American families. Gender bias in resource allocation has been studied widely in the context of developing countries and studies have found marked differences in resource allocation, depending on child gender; and these differences vary with the gender of the parent (Thomas; 1990, 1994). To date, however, empirical evidence for developed countries, such as the United States, has been scarce. We explore the differences in preventive care consumption between sons and daughters and test the hypothesis that parents exhibit gender preference in allocation of resources under their control.

The results suggest the ongoing importance of women's traditional gender role as caregiver and are in line with the previous findings that consumption of goods enhancing household well-being is higher in families in which women have greater control over economic resources. The analysis reveals that the mother's non-labor income offers a greater positive impact than the father's non-labor income on children's preventive care demand. We find no evidence supportive of gender bias or parental gender preference in child health investment. The wife's non-labor income and education have greater positive impacts on her preventive care visits than do her husband's income and education. Spouses' incomes are not significant determinants of the husband's preventive care demand and the income pooling cannot be rejected for him. The wife's education has a greater positive impact on the husband's preventive care demand than does his own education level.

In the next section, we provide a brief review of the related literature. Section 3 presents the theoretical framework based on the unitary model and an alternative collective model (namely the Nash bargaining model) and their implications. Empirical specification and a description of the data are given in section 4 and section 5 presents the empirical results. Section 6 concludes.

4.2 RELATED LITERATURE

A considerable amount of empirical work tests the resource-pooling hypothesis of the unitary model by using different indicators of controlling power over resources and different outcomes of interest. In this review, we focus on the studies examining the relationships between female resource control and subsequent household well-being. Several previous studies focused on the household expenditures on different types of goods that presumably lead to favorable or unfavorable outcomes concerning household well-being. Haddad and Hoddinott (1994, 1995), using data from Cote d'Ivoire, find that the share of woman's cash income increases the budget share of food and reduces the budget share of alcohol and cigarettes. Beegle et al. (2001) examine whether a woman's power, relative to her husband's, affects decisions about use of prenatal and delivery care in Indonesia. Contrasted with women with no assets at all, a woman holding some share of household assets influences reproductive health decisions. Although there has been considerably less empirical work in the context of the developed country, Phipps and Burton (1998), using Canadian data, reject the income-pooling hypothesis for seven out of twelve expenditure categories. They discerned that an increase in the wife's income, relative to her husband's income, correlates with greater expenditures on childcare, woman's clothing, and with reduced expenditures on alcohol and tobacco. Lundberg, Pollak, and Wales (1997) found that when the United Kingdom began to give governmental child benefit to the mother, rather than to the father as previously administered, household expenditures on children's clothing increased. These empirical findings have been interpreted to be evidence that men and women have different preferences regarding household consumption priorities.

A key area of inquiry focuses on household decision-making for investments in children. Several studies explore the relationship between female decision-making power and child outcomes in developing countries. These studies use either woman's income share or education to measure her power within her family unit. It has been shown that the caloric intake of children is affected positively by female wage and negatively affected by male wage (Senauer, Garcia & Jacinto, 1988). In addition, children's height for their age increases with additional female education, while male education demonstrates both positive and negative effects in different studies (Senauer & Garcia, 1991; Handa, 1999). Thomas (1990) uses data from a Brazilian sample and finds that non-labor incomes of husband and wife have different effects on child health. Strikingly, the affect of mother's non-labor income on child survival probabilities is almost twenty times that of father's. Thomas et al. (2002) uses the value of resources that husbands and wives in Indonesia brought to the marriage to measure their relative control in the union itself. He finds that a child's health is influenced by the relative asset positions of parents at the time of marriage, even after controlling current household resources.

Additionally, several studies show that parents do not have identical preferences towards daughters and sons in certain parts of the developing world. Duflo (2000) finds that pension received by women had a large effect on the anthropometric outcomes of girls, but little effect on that of boys. However, there was no observation of similar effects for pensions received by men. In much of South and East Asia, girls receive less nutritious foods and less education and health care (Behrman, 1992). In contrast, for the U.S., there currently exists limited evidence that parents purchase goods and services related to human-capital formation according to their child's sex. Taubman (1991) concludes that there is little evidence of differential treatment of children by gender in bequests, in transfers, and in educational attainment. There is, however, evidence that the time parents spend with their children differ by the child's sex. Time allocation data from the United States show that men spend more time with sons and women spend more time with daughters (Bryant and Zick, 1996; Yeung et al., 2001).

This paper extends the literature in several ways. First, we use data on individual specific consumption of preventive medical care. Observing individual specific demands allows the examination of how consumption patterns evolve for each family member as the balance of power shifts within the household. We estimate preventive care demand functions for wives, husbands and children. Second, previous studies inferred the allocation of resources by focusing on a set of allocation outcomes. For example, they examine differential impact of parental incomes on child health and infer the families' child health resources allocations. By examining medical care consumption of family members, we present direct evidence (direct in that we examine resource allocations rather than merely the outcomes themselves) on the process of allocation that may lead to better outcomes. Third, most studies tested the unitary model using data from developing countries; we use U.S. data in this paper. Last, we provide empirical evidence for gender discrimination in child health investment in American families.

4.3 Theoretical Framework

The unitary model assumes that the household has a common set of preferences and maximizes household utility subject to a family budget constraint. Husband's and wife's preferences over health capital investments of family members (husband, wife and children) and a composite good Z are represented by

$$U = U(M_h, M_w, M_c, Z) \tag{1}$$

where U is the household utility function and M_j (j=h,w,c) denotes preventive care demands of husband (h), wife (w) and child (c). The family budget constraint is given by

$$P_m(M_h + M_w + M_c) + P_z Z = A + \sum w_i T \tag{2}$$

where P_m and P_z are the prices of preventive care services and composite good, respectively; A is the total family non-labor income; w_i is the wage rate of the ith individual (i=husband, wife) and T is the total time available. Maximization of the utility function (1) subject to the family budget constraint (2) yields a set of demand functions

$$j = f_j(P_m, P_z, w_h, w_w, A), \qquad j = M_h, M_w, M_w, Z \tag{3}$$

Since the model assumes that husband and wife pool their resources, the demand for a specific good is expressed as a function of total non-labor income, not the individual non-labor incomes.

The collective models of household introduce divergent preferences within the family and they analyze the household behavior within the frameworks of cooperative and non-cooperative game theory. Empirical testing of many collective models has been sparse, given their complexity; many of them do not imply a unique equilibrium. Among the collective models, the Nash bargaining approach has been the most widely used in empirical studies. Since one of the implications of this approach can be empirically tested, we briefly present the Nash bargaining set-up as an example of an alternative framework to the unitary approach.

The model assumes that the husband and wife jointly allocate resources according to a two person Nash cooperative game. The individual's utility function is given by

$$U_i = U_i(M_h, M_w, M_c, Z), \qquad i = h, w \tag{4}$$

where U_i is the utility of individual i (i=h,w). The individual budget constraint is

$$P_m(M_h + M_w + M_c) + P_z Z = A_i + w_i T \tag{5}$$

where A_i is the respective non-labor incomes of the husband and wife. The implied indirect utility function is given by

$$V_i = V_i(P_m, P_z, w_i, A_i) \tag{6}$$

where V_i serves as a threat point (i.e., the maximum level of utility attainable outside the marriage) for husband and wife. The Nash-bargained solution requires that the husband and wife choose commodities in the utility function to maximize the product of the gains from marriage, i.e., to maximize

$$G = \prod [U_i(M_h, M_w, M_c, Z) - V_i(P_m, P_z, w_i, A_i)] \tag{7}$$

subject to the family budget constraint:

$$P_m(M_h + M_w + M_c) + P_z Z = A_h + A_w + (w_h + w_w)T \tag{8}$$

The solution to this household optimization problem yields the demand functions

$$j = f_j(P_m, P_z, w_h, w_w, A_h, A_w) \qquad j = M_h, M_w, M_w, Z \tag{9}$$

These demand functions include individual specific non-labor incomes, as compared to the demand functions in (3), in which non-labor incomes are pooled. This distinction suggests a simple test of the unitary model against the alternative Nash bargaining model. One can test whether the non-labor income of husband and wife affect family consumption decisions differently (bargaining approach) or if they have the same effect (unitary approach). Since the implication of relative control over resources affecting household demand is common to all collective models, not just exclusively to the Nash bargaining model, the rejection of the equality of these effects has implication for the entire class of collective models. Thus, our work tests the unitary model against a broad class of alternatives.

4.4 EMPIRICAL SPECIFICATION AND DATA

The implication that non-labor incomes should enter the demand equations in the same way is a key feature of the unitary model not shared by collective models. We estimate reduced form preventive care demand functions for children, husbands and wives separately. The preventive care demand equation that is estimated for children is

M_c= a + β_1 Mother's non-labor income + β_2 Father's non-labor income + β_3 Mother's years of education + β_4 Father's years of education + φX_1 + ε

and the equations that are estimated separately for husbands and wives are

$M_s = b + \theta_1$ Own non-labor income + θ_2 Spouse's non-labor income + θ_3 Own years of education + θ_4 Spouse's years of education + $\psi X_2 + \upsilon$,

where M is a measure of preventive care utilization, and X_1, X_2 are vectors of individual and family characteristics that affect demand.

If family members pool their resources and allocate the total to maximize a single objective function, then only the total household income will affect demands. The unitary model implies that non-labor income has an identical effect on household demands, regardless of the source of that non-labor income. Similarly, if education is a good indicator of decision-making power, a finding of unequal effects of spousal education implies that spouses do not have common preferences towards preventive care use; the spouse having more controlling power dominates in the determination of individual preventive care demands.

The Medical Expenditure Panel Survey (MEPS) provided the data used in this study. The MEPS, which began in 1996, is an annual, nationally representative survey of households that interviews over 20,000 people. The MEPS consists of a family of three interrelated surveys: the Household Component, the Medical Provider Component, and the Insurance Component. Of those three elements, we utilize data from the MEPS Household Component for this study. The set of households selected for the Household Component is a sub-sample of those participating in the National Health Interview Survey. At each interview, the questionnaire accrues information about each household member, and one designated household respondent reports all MEPS data. The survey consists of an overlapping panel design that, each year, initiates a new panel followed for two years through five in-person interviews.

To allow adequate sample sizes for examination of groups of particular policy interest, the MEPS over-samples certain populations. Furthermore, the sample design of the survey includes stratification, clustering and multiple stages of selection. By applying MEPS survey weights to produce estimates and by using an appropriate technique to derive standard errors associated with the weighted estimates, analysts need to account for the complex survey design.

A series of calendar year-specific MEPS data files are produced annually. These files contain data from the second year of a continuing panel in conjunction with the first year of a new panel. The person identifier remains the same for a person for his/her entire duration in the survey. The survey provides detailed information on demographic characteristics, health status, medical services utilization, health insurance coverage, income and employment for every member of the household. Regarding medical care services usage, the survey collects data about all hospital (emergency room, inpatient and outpatient events), physician services, home health care and prescribed medicines.

For the purposes of this study, we pool data from the 2002, 2003 and 2004 surveys. The MEPS assigns family identification numbers to individuals, allowing construction of family units. The MEPS definition of family includes all persons living in the same household related by blood, marriage and adoption or self-identified as a single unit; hence, their definition includes non-married partners, foster children, in-laws and other relatives. Because such extended families are less likely to make family health care decisions together, instead of this broad definition, we construct the family units using the health insurance eligibility units defined by the MEPS as corresponding to groups of people who would typically have family coverage under a common health insurance policy.

We initially restrict the sample to families in which a husband and a wife both are present, and then, to focus on spousal decision making, further narrow the sample to households with no “third-party” adults. We examine preventive care utilization of children, wives and husbands separately. The sample of children includes all children below 17 years of age whose parents both are present in the household. A total of 6,054 children are in the sample. The sample of wives includes married women 18 to 55 years of age whose spouses are present in the family unit. Similarly, the sample of husbands includes all married men 18 to 55 years of age whose spouses are present in the family unit. This definition gives a sample of 4,137 husbands and 4,523 wives.

To measure decision-making power of the spouses, we utilize spousal non-labor incomes and education attainments. In this study, non-labor income refers to income from physical and financial assets, public assistance and all other non-wage income. In order to qualify the robustness of non-labor income effects, we use asset incomes of spouses to measure their control over resources. Asset income refers to the interest, dividend, trust and rent income.

To measure family members’ preventive care demand, we use annual number of preventive care visits. The MEPS provides detailed information on each medical visit of family members during the year. Preventive visits are identified using visit-level information on office-based visits. The MEPS asks the primary reason for the visit and categorize each visit based on the type of medical service received. For this analysis, visits are coded as preventive if the primary reason given was “general check-up”, “well child exam” or “immunizations or shots”.

Following the literature on medical care demand, the child health care demand function includes age, gender, race/ethnicity, family size, health status, insurance status and an indicator for the survey respondent as controls. We use insurance status as a proxy for the medical care prices faced by family members. In this specification, a significant coefficient for the gender indicator suggests differential health investment based on child gender. In order to study parental gender preference in child health investments, we estimate an additional equation in which we include interaction variables obtained by multiplying parental resource control variables with child gender.

For wives' and husbands' health care, spousal power measures aside, demand equations include controls for age, race/ethnicity, family size, health status, insurance status and an indicator for the survey respondent. We use reported health status indicators for poor/fair health, excluding the category of good/very good/excellent health. We also include an indicator that takes the value one if the person has any health limitations (e.g. physical, social or cognitive limitations). A health insurance coverage indicator is used to control for insurance status. Table 4.1 presents descriptive statistics of the variables used in the study.

The econometric model for health care utilization needs to account for the nature of the dependent variable, which is a non-negative integer valued count. A natural starting point for the analysis is the poisson regression model, which assumes that the conditional mean is equal to conditional variance. Since this assumption is not appropriate in our case where the data show over-dispersion, we adopt the negative binomial regression to model preventive care demand. We adjust all the estimates in this study for the survey sample design using the survey weights. The standard errors for the weighted estimates also are adjusted for the complex survey design.

4.5 EMPIRICAL RESULTS

4.5.1 Children's Preventive Care Demand

Table 4.2 presents the negative binomial regression results for the number of preventive care visits of children. The main predictors are the parents' non-labor incomes and education attainments. Higher values presumably correspond to greater decision-making power. We test whether parental resources have the same effect on children's preventive care demand. Table 4.2 also presents the Wald test results for the equality of the effects of parental resources.

The results show that the mother's non-labor income significantly increases child preventive care demand; while the effect of the father's non-labor income is statistically insignificant. We reject the unitary model. The Wald test statistic implies that the effects of the mother and the father's non-labor incomes are not the same (p-value=0.028). Income in the hands of the mother leads to greater preventive care utilization for children. This result supports the conclusion of the existing literature: fathers and mothers have different preferences for their children's health. Mothers appear to be more interested in their children's welfare than are fathers; consequently, child health investment increases in direct correlation to increases in the mother's available resources. In addition, parents' education levels have a positive and significant effect. While maternal education is more influential than paternal education, the difference lacks statistical significance (p=0.322).

The coefficient estimates for the other covariates are plausible. For example, children in fair/poor health and those with a health limitation use more preventive medical care, and uninsured children have fewer visits than insured children do. Preventive care demand also decreases as children age. Family size has a negative effect on children's health care demand. Hispanic children and black children have less preventive visits than do white children. Last, if the mother is the survey respondent, children are reported to utilize more preventive care.

4.5.1.1 Gender differences

The results do not suggest differential health investment based on child gender in American families. In Table 4.2, the indicator for male child is statistically insignificant, which is in line with other studies that find no evidence of differential human capital investment according to child sex in the United States. One can conjecture that different race/ethnicity groups have different social norms, so that some may place different values on boys and girls. To test this, we estimate the regression equation separately for each race/ethnicity group. In each case, the indicator for male child is statistically insignificant[4].

Although we do not see gender bias overall, parents may exert gender preference in allocating the resources under their control. There is evidence for developing countries that mothers prefer to devote resources to improving the outcomes for their daughters, as do fathers to sons. To test for this, we include in the basic regression the interaction variables obtained by multiplying parental non-labor incomes with the gender indicator. To test for differential education effects, we also interact the male indicator with parental education. Significantly positive coefficients for these variables would suggest parents' son preference in health investment. Table 4.3 reports the estimation results. None of the interaction variables is statistically significant, suggesting that parental resources have the same impact on both boys and girls. Repeating this analysis for each race/ethnicity

[4] For the sake of brevity, we do not include these results.

groups separately does not change the results. For brevity, these results also are excluded. Contrary to the findings of most studies using data from developing countries, we find that parents do not to exhibit gender preference in child health investment.

4.5.1.2 Robustness of Estimated Parental Non-labor Income and Education Effects

One can argue that differences between people having no non-labor income and those having positive non-labor income drive the estimated differential effects of parental incomes. To deal with this conjecture, we re-estimate the basic regression equation constraining the sample to include only children with parents reporting positive non-labor income. The findings are presented in Table 4.4. Again, income-pooling is rejected at the 5 percent level; the mother's resources have greater impact on child preventive care than do the father's resources. Persistence of the differential effect on this sub-sample suggests that the issue of differences between people without non-labor income and those with positive non-labor is not relevant.

As a second robustness check, we replace parents' non-labor incomes with their asset incomes. Table 4.5 presents the results. The mother's asset income has a positive effect on child preventive care use and is statistically significant at the 5 percent. The father's asset income is not statistically significant. The null hypothesis of equal parental income effects is rejected at the 5 percent level. Overall results for children indicate that non-labor income in control of the mothers has a larger impact on child health investment than that of the fathers.

A key area of inquiry in the literature has been the socio-economic determinants of child health care utilization. The findings in this paper suggest that a new factor, distribution of income between spouses, needs to be taken into account in studying family decisions regarding the demand for child health care. Children appear to fare better in receiving preventive health care when the mother's exogenous income increases relative to that of the father.

4.5.2 Wives' and Husbands' Preventive Care Demand

The negative binomial regression results for the number of annual preventive care visits of wives and husbands are presented in Table 4.6. The key variables are spousal non-labor incomes and education attainments. Higher values presumably correspond to greater control over resources. The Wald test results for whether spousal resources have the same effect on wives and husbands' preventive care utilization also are presented in Table 4.6.

The first column of Table 4.6 shows the results for the wives. The wife's own non-labor income significantly increases her own preventive care use. Her husband's non-labor income, on the other hand, is not a significant determinant. This differential income effect is significant at the 5 percent level. We reject the income-pooling hypothesis of the unitary model.

Both the wife's education and her husband's education have a significantly positive effect on the wife's number of preventive visits. The estimated coefficient of the wife's education is larger than the coefficient of her spouse's education. This differential education effect is significant at the 5 percent level. If education is an appropriate proxy for spousal resource controlling power, this finding implies that as the wife gains more power relative to her husband, she utilizes more preventive care.

The second column of Table 4.6 presents the results for the husband's preventive care demand. In this case, neither his nor his wife's non-labor income is significant. We cannot reject the unitary model (p-value=0.270). The identity of income receiver does not seem to matter in the determination of the husband's preventive care consumption.

Regarding the education effects, both the husband's education and his wife's education have a significantly positive effect on the husband's preventive care visits. It is especially noteworthy that the wife's education has a greater influence on her husband's preventive care demand than the husband's education does on his own care. This differential education effect is significant at the 5 percent level. The human capital characteristics of women seem to be more important in determining their husband's preventive care utilization than does their income levels.

The coefficient estimates for the other covariates are as expected. Adults in fair/poor health and those with a health limitation use more preventive medical care. Uninsured people have fewer visits than do the insured and preventive care demand increases with age. Family size has a negative effect on adults' preventive care demand. Hispanic people and black people have less preventive visits than do white people.

4.5.2.1 Robustness of Estimated Spousal Non-labor Income and Education Effects

We repeat the previously discussed robustness checks for spouses' preventive care demand. Table 4.7 shows the estimation results when the sample is restricted to include only those spouses with positive non-labor incomes. The findings are robust to this specification; significantly different income and education effects are found. Specifically, the wife's non-labor income significantly increases her own preventive care demand, whereas her husband's non-labor income is not significant. Both the wife's and the husband's education have a significantly positive effect; nevertheless, the influence of the wife's education level is larger. The second column of Table 4.7 shows that the husband's non-labor income has a negative effect on his preventive care use, which is marginally significant at the 10 percent level. The wife's non-labor income is statistically insignificant. Both spouses' education increases the husband's preventive visits; the wife's education, however, has a significantly larger effect. Robustness of the results on this sub-sample suggests that the findings are not driven by differences between people with no non-labor income and those with positive non-labor income.

Table 4.8 reports the results when spouses' asset incomes replace their non-labor incomes. In this case, all of the income variables are statistically insignificant: asset income is not a significant determinant of spouses' preventive care demand. However, the results for education variables are, again, robust. The husband's and the wife's preventive care demands increase with both spouses' education level, but the wife's education is significantly more influential for both spouses' preventive visits.

Overall results for spouses suggest that non-labor income in control of the wives has greater impact on their own health investment but no effect on their husbands' health investment. The wives' education has a larger impact both on their own and on their husbands' health investment.

4.6 SUMMARY AND CONCLUSION

This paper examines the impact of changes in the resource controlling power of spouses, measured by their non-labor incomes and education attainments, on family members' preventive care consumption. Since the unitary (i.e., common preference) model assumes no impact, the statistical significance of this impact or otherwise constitutes a test of the unitary versus collective model of the household. There is recognition that giving control over money to women is more effective in increasing household well-being because women have stronger preferences for it. We focus on the role of the wife's resources in determining the preventive care demands of individual members of the family. Since health care utilization is often within the women's realm of responsibility when adhering to traditional gender roles, women may have stronger preferences for preventive medical care. An implication that emerges from a collective model is that when women have relatively more resources, individual family members utilize more preventive care.

Our work contributes to the literature in two broad ways. First, it used data from a developing country; most of the literature has worked with data from the developing world. Second, we employ individual specific medical care consumption data. Observing individual consumption allows the examination of how resource allocation changes for each family member as the balance of income controlling power changes within the family. We estimate separate demand equations for children, wives and husbands.

The findings suggest that the resources of the mother have a greater positive impact on child preventive care demand than does the resources of the father. This supports the view of the existing literature that women are more dedicated to using resources for child benefit than are men, and that, accordingly, children fare better as their mother's control over family resources increases. In contrast to most studies from developing countries, we find that US families do not exhibit differential health care demand based on child gender. This is in line with other studies that find no evidence of differential human capital investment according to child sex in the United States. Furthermore, parents do not exhibit gender preference in the allocation of resources under their control.

Increasing preventive care utilization among children has long been a policy target. Understanding the determinants of preventive care utilization is essential for achieving this target. These findings suggest a new factor that is important in the determination of children's health care utilization: distribution of income between spouses. Children fare better in receiving preventive health care when the mothers' income increases relative to that of the fathers.

The estimation results for spouses' health care suggest that the wife's non-labor income and her education level have a greater positive impact than those of her husband do on her own preventive care utilization. The unitary model is rejected. For husbands, we cannot reject the unitary model: non-labor incomes of both spouses have the same impact on the husbands' preventive care visits. These effects, however, are not significant. It is also noteworthy that the wife's education has a greater positive impact on the husband's preventive care demand than does his own education level. The wife's human capital characteristics seem to be more important in determining her husband's preventive care demand than does her income level.

The overall results indicate the ongoing importance of the traditional gender role of women as the primary caregiver and are in line with the findings of previous studies that consumption of goods that enhance household well-being is higher in families in which the wife has greater control over economic resources. This has implications for the design of policies designed to transfer resources to households: the identity of the transfer recipient does affect the ultimate outcome of the intervention. In this particular case, a transfer program that increases women's exogenous income leads to greater preventive care utilization for both children and women.

Tables and Figures

Table 2.1 Sample Statistics

Table 2.1 Sample Statistics

	Mean	Standard Deviation
Compliance ratio	65.987	(32.749)
Number of well-child care visits	3.220	(1.684)
Fair or poor health in year 2	0.019	(0.136)
History of a serious illness in year 2	0.221	(0.414)
Health limitation in year 2	0.024	(0.154)
Fair or poor health in year 1	0.022	(0.147)
History of a serious illness in year 1	0.203	(0.4020
Health limitation in year 1	0.041	(0.198)
Number of reported health conditions	2.732	(2.370)
Age in months	19.206	(9.390)
Male	0.509	(0.500)
Hispanic	0.216	(0.412)
Black	0.145	(0.352)
Family income	48,640	(41,760)
Family size	4.049	(1.413)
Mother employed	0.640	(0.480)
Mother's years of education	12.804	(3.006)
Mother's age	29.276	(6.253)
Father not present	0.304	(0.460)
MSA	0.824	(0.381)
Any private insurance	0.663	(0.303)
Only public insurance[a]	0.236	(0.404)
Northeast	0.171	(0.377)
Midwest	0.218	(0.413)
West[b]	0.240	(0.427)
Number of observations	2,958	

[a]Uninsured is the excluded category.
[b]South is the excluded category.

Figure 2.1 Histogram of compliance ratio

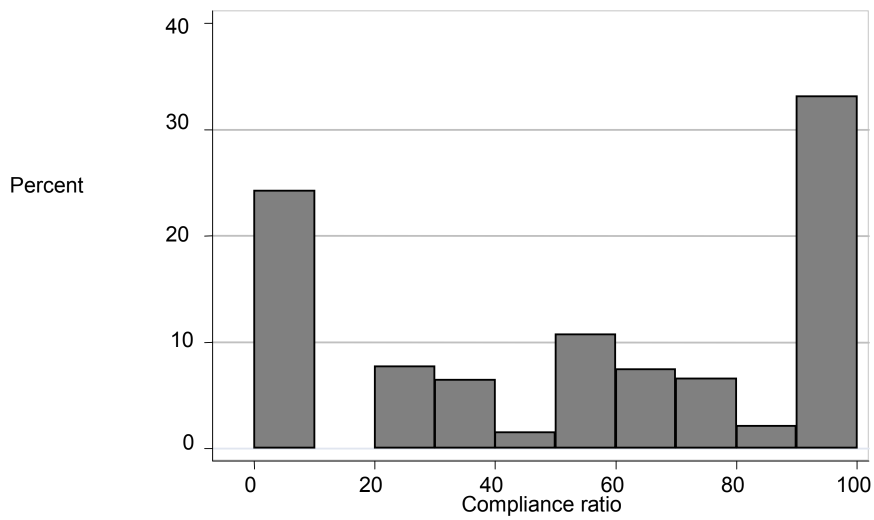

Figure 2.1 Histogram of compliance ratio

Table 2.2 Determinants of Fair or Poor Health in Year 2

Table 2.2 Determinants of Fair or Poor Health in Year 2

	Ordinary Least Squares Estimates	
Compliance ratio	-0.0001**	(5.1e-05)
Fair or poor health in year 1	0.2681***	(0.0623)
Number of reported health conditions	0.0032***	(0.0010)
Age in months	-0.0001	(0.0003)
Male	-0.0009	(0.0039)
Hispanic	0.0177***	(0.0066)
Black	0.0037	(0.0072)
Family income/10,000	-0.0007	(0.0006)
Family size	0.0006	(0.0019)
Mother employed	-0.0039	(0.0047)
Mother's years of education	-0.0010	(0.0009)
Mother's age	0.0006	(0.0005)
Father not present	-0.0015	(0.0081)
MSA	-0.0098	(0.0071)
Constant	0.0178	(0.0231)
Number of observations		2,958
R^2		0.101
F- test statistic for test of relevance of instruments		16.45***
		[0.00]
Hansen J statistic for test of exclusion restrictions		2.17
		[0.82]
Rivers-Vuong test statistic for test of exogeneity		0.18
		[0.67]

Notes: Robust standard errors are in parenthesis. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All of the explanatory variables are measured in the first-year survey. Numbers in brackets are the p values for specification test statistics.

Table 2.3 Determinants of Having a History of a Serious Illness in Year 2

Table 2.3 Determinants of Having a History of a Serious Illness in Year 2

	Ordinary Least Squares Estimates		Two-Stage Least Squares Estimates	
Compliance ratio	-0.0003*	(0.0002)	-0.0027**	(0.0011)
History of a serious illness in year 1	0.2733***	(0.0330)	0.2536***	(0.0343)
Number of reported health conditions	0.0139**	(0.0058)	0.0173***	(0.0061)
Age in months	0.0007	(0.0014)	-0.0015	(0.0023)
Male	0.0156	(0.0231)	0.0070	(0.0259)
Hispanic	-0.0307	(0.0348)	-0.0614	(0.0455)
Black	-0.0104	(0.0334)	-0.0334	(0.0396)
Family income/10,000	0.0014	(0.0030)	0.0014	(0.0033)
Family size	0.0005	(0.0078)	-0.0091	(0.0100)
Mother employed	0.0184	(0.0234)	0.0132	(0.0242)
Mother's years of education	-0.0123***	(0.0043)	-0.0105**	(0.0046)
Mother's age	0.0006	(0.0023)	0.0024	(0.0026)
Father not present	0.0500*	(0.0293)	0.0495*	(0.0300)
MSA	0.0407	(0.0312)	0.0430	(0.0338)
Constant	0.2100**	(0.0822)	0.3767	(0.1405)
Number of observations		1,729	1,729	
R^2		0.097	0.106	
F- test statistic for test of relevance of instruments			12.60*** [0.00]	
Hansen J statistic for test of exclusion restrictions			8.24 [0.14]	
Rivers-Vuong test statistic for test of exogeneity			6.23** [0.03]	

Notes: Robust standard errors are in parenthesis. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All of the explanatory variables are measured in the first-year survey. Numbers in brackets are the p values for specification test statistics.

Table 2.4 Determinants of Having a Health Limitation in Year 2

Table 2.4 Determinants of Having a Health Limitation in Year 2

	Ordinary Least Squares Estimates	
Compliance ratio	-0.0002**	(8.0e-05)
Health limitation in year 1	0.0984***	(0.0343)
Number of reported health conditions	0.0060***	(0.0018)
Age in months	-0.0010**	(0.0005)
Male	0.0031	(0.0071)
Hispanic	-0.0029	(0.0072)
Black	0.0057	(0.0103)
Family income/10,000	-0.0006	(0.0006)
Family size	0.0022	(0.0021)
Mother employed	-0.0075	(0.0076)
Mother's years of education	-0.0002	(0.0010)
Mother's age	0.0001	(0.0005)
Father not present	0.0025	(0.0070)
MSA	0.0052	(0.0076)
Constant	0.0221	(0.0241)
Number of observations		2,313
R^2		0.037
F- test statistic for test of relevance of instruments		13.77***
		[0.00]
Hansen J statistic for test of exclusion restrictions		6.32
		[0.20]
Rivers-Vuong test statistic for test of exogeneity		0.37
		[0.56]

Notes: Robust standard errors are in parenthesis. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All of the explanatory variables are measured in the first-year survey. Numbers in brackets are the p values for specification test statistics.

Table 2.5 Effect of Number of WCC Visits on Health Outcomes

Table 2.5 Effect of Number of WCC Visits on Health Outcomes

	Fair or poor health in year 2		History of a serious illness in year 2		Health limitation in year 2	
Number of WCC visits	-0.0027	(0.0017)	-0.0085	(0.0075)	-0.0047*	(0.0025)
Fair or poor health in year 1	0.2681***	(0.0624)	-	-	-	-
History of a serious illness in year 1	-	-	0.2840***	(0.0326)	-	-
Health limitation in year 1	-	-	-	-	0.0989***	(0.0342)
Number of reported health conditions	0.0144***	(0.0041)	0.0796**	(0.0393)	0.0062***	(0.0018)
Age in months	-0.0003	(0.0004)	0.0006	(0.0015)	-0.0012**	(0.0005)
Male	-0.0006	(0.0039)	0.0187	(0.0229)	0.0032	(0.0071)
Hispanic	0.0173***	(0.0066)	-0.0314	(0.0347)	-0.0035	(0.0073)
Black	0.0036	(0.0072)	-0.0158	(0.0329)	0.0055	(0.0102)
Family income/10,000	-0.0008	(0.0006)	0.0019	(0.0030)	-0.0006	(0.0006)
Family size	0.0007	(0.0019)	0.0004	(0.0079)	0.0021	(0.0021)
Mother employed	-0.0041	(0.0047)	0.0181	(0.0232)	-0.0078	(0.0076)
Mother’s years of education	-0.0010	(0.0009)	-0.0120***	(0.0043)	-0.0002	(0.0010)
Mother’s age	0.0006	(0.0005)	0.0006	(0.0023)	0.0001	(0.0005)
Father not present	-0.0016	(0.0081)	0.0520*	(0.0298)	0.0023	(0.0071)
MSA	-0.0089	(0.0066)	0.0425	(0.0315)	0.0055	(0.0076)
Constant	0.2835***	(0.0681)	0.4526***	(0.1020)	0.0253	(0.0249)
Number of Observations	2,958		1,729		2,313	
R^2	0.133		0.126		0.040	

Notes: Regressions are estimated by Ordinary Least Squares. Robust standard errors are in parenthesis. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All of the explanatory variables are measured in the first-year survey.

Table 3.1 Sample Statistics

Table 3.1 Sample Statistics

	Mean	Standard Deviation
Had preventive visit	0.477	(0.500)
Had curative visit	0.550	(0.498)
Number of preventive visits	0.968	(1.926)
Number of curative visits	2.009	(5.524)
Mother working	0.643	(0.474)
Mother's hours of work per week[1]	35.040	(11.186)
Mother's self-employment hours per week[2]	30.801	(18.546)
Mother's wage employment hours per week[3]	35.610	(10.856)
Mother's earnings per week[4]	431.514	(325.71)
Other household income per week	1437.421	(1256)
Other adults present in the household	0.770	(0.421)
Mother's age	33.889	(6.916)
Mother's education	12.914	(2.767)
Father present	0.731	(0.444)
Child age <2	0.211	(0.408)
Child age 3-5[5]	0.243	(0.429)
Male	0.512	(0.500)
Hispanic	0.167	(0.373)
Black[6]	0.150	(0.357)
Fair or poor health	0.014	(0.116)
Good health[7]	0.118	(0.323)
Health limitation	0.024	(0.154)
Family size	7.575	(3.130)
Northeast	0.185	(0.388)
Midwest	0.241	(0.428)
West[8]	0.244	(0.430)
MSA	0.822	(0.382)
Mother is the survey respondent	0.823	(0.382)
Insured	0.880	(0.324)
Number of observations	19,390	

[1] Working mothers only.
[2] Self-employed mothers only.
[3] Wage-employed mothers only.
[4] Working mothers only.
[5] Excluded category is child age 6-12.
[6] Excluded category is white.
[7] Excluded category is very good/excellent health.
[8] Excluded category is south.

Table 3.2 Maternal Labor Market Participation (%)

Table 3.2 Maternal Labor Market Participation (%)

	Mother not working	Mother wage-employed	Mother self-employed
All children	36	51	13
Children<2	37	53	10
Children age 3-5	38	49	13
Children age 6-12	34	53	13

Table 3.3 Impact of Maternal Work on Probability of Any Preventive Visit

Table 3.3 Impact of Maternal Work on Probability of Any Preventive Visit

	(1)		**(2)**		**(3)**	
Mother's hours of work/week	-0.0029***	(0.0006)	-	-	-0.0033***	(0.0008)
Mother's earnings/week	0.0001***	(2.0e-05)	0.0001***	(2.0e-05)	0.0001***	(2.0e-05)
Mother working	0.0163	(0.0223)	0.0073	(0.0208)	0.0159	(0.0220)
Mother's self-employment hours/week	-	-	0.0001	(0.0005)	-	-
Mother's wage-employment hours/week	-	-	-0.0029***	(0.0006)	-	-
Mother's hours x child age <2	-	-	-	-	0.0017**	(0.0007)
Mother's hours x child age 3-5	-	-	-	-	0.0004	(0.0008)
Mother's earnings x child age <2	-	-	-	-	-0.0001**	(4.0e-05)
Mother's earnings x child age 3-5	-	-	-	-	-0.00002	(0.0001)
Other household income/week	0.00002***	(5.2e-06)	0.00002***	(5.3e-06)	0.00002***	(5.3e-06)
Number of observations	19,390		19,390		19,390	
F statistic for test of equal income effects	18.99*** [0.00]		14.60*** [0.00]		11.63*** [0.00]	

Notes: The Ordinary Least Squares estimation results for the probability of at least one preventive visit during the year. Robust standard errors are in parentheses.*** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All specifications also include mother's age and education, family size, indicator variables for child's age (less than 2, 3 to 5 and 6-12-excluded-), sex (male), race/ethnicity (Hispanic, black and white -excluded-), health status (fair or poor health, good health and excellent/very good health-excluded-), insurance status (insured), census region (Northeast, Midwest, West, South-excluded), presence of father in the household, presence of other adults in the household, having a health limitation, residence in a metropolitan statistical area, whether the mother is the survey respondent and a constant. The p value for the F statistic is given in the bracket.

Table 3.4 Impact of Maternal Work on Number of Preventive Visits

Table 3.4 Impact of Maternal Work on Number of Preventive Visits

	(1)		**(2)**		**(3)**	
Mother's hours of work/week	-0.0068***	(0.0017)	-	-	-0.0091***	(0.0020)
Mother's earnings /week	0.0003***	(0.0001)	0.0003***	(0.0001)	0.0004***	(0.0001)
Mother working	0.0705	(0.0667)	0.0500	(0.0682)	0.0692	(0.0661)
Mother's self-employment hours/week	-	-	0.0009	(0.0016)	-	-
Mother's wage-employment hours/week	-	-	-0.0063***	(0.0020)	-	-
Mother's hours x child age <2	-	-	-	-	0.0069***	(0.0020)
Mother's hours x child age 3-5	-	-	-	-	0.0027**	(0.0011)
Mother's earnings x child age <2	-	-	-	-	-0.0003***	(0.0001)
Mother's earnings x child age 3-5	-	-	-	-	-0.0002**	(0.0001)
Other household income/week	0.00004***	(1.4e-05)	0.00004***	(1.4e-05)	0.00004***	(1.4e-05)
Number of observations	19,390		19,390		19,390	
F statistic for test of equal income effects	11.78*** [0.00]		8.65*** [0.00]		13.63*** [0.00]	

Notes: The Negative Binomial Regression results for the number of annual preventive care visits of children. Robust standard errors are in parentheses.*** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All specifications also include mother's age and education, family size, indicator variables for child's age (less than 2, 3 to 5 and 6-12-excluded-), sex (male), race/ethnicity (Hispanic, black and white -excluded-), health status (fair or poor health, good health and excellent/very good health-excluded-), insurance status (insured), census region (Northeast, Midwest, West, South-excluded), presence of father in the household, presence of other adults in the household, having a health limitation, residence in a metropolitan statistical area, whether the mother is the survey respondent and a constant. The p value for the F statistic is given in the bracket.

Table 3.5 Impact of Maternal Work on Probability of Any Curative Visit

Table 3.5 Impact of Maternal Work on Probability of Any Curative Visit

	(1)		(2)		(3)	
Mother's hours of work/week	-0.0020***	(0.0005)	-	-	-0.0020***	(0.0006)
Mother's earnings /week	0.0001***	(0.0000)	0.0001**	(0.0000)	0.00003	(2.4e-05)
Mother working	0.0558***	(0.0190)	0.0536***	(0.0190)	0.0543***	(0.0190)
Mother's self-employment hours/week	-	-	-0.0009	(0.0006)	-	-
Mother's wage-employment hours/week	-	-	-0.0018***	(0.0005)	-	-
Mother's hours x child age <2	-	-	-	-	0.0007	(0.0008)
Mother's hours x child age 3-5	-	-	-	-	-0.0009	(0.0008)
Mother's earnings x child age <2	-	-	-	-	0.0001	(0.0001)
Mother's earnings x child age 3-5	-	-	-	-	0.0002	(0.0003)
Other household income/week	0.00002***	(4.8e-06)	0.00002***	(4.9e-06)	0.00002***	(4.9e-06)
Number of observations	19,390		19,390		19,390	
F statistic for test of equal income effects	2.69 [0.11]		1.90 [0.17]		1.23 [0.13]	

Notes: The Ordinary Least Squares estimation results for the probability of at least one curative visit during the year. Robust standard errors are in parentheses.*** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All specifications also include mother's age and education, family size, indicator variables for child's age (less than 2, 3 to 5 and 6-12-excluded-), sex (male), race/ethnicity (Hispanic, black and white -excluded-), health status (fair or poor health, good health and excellent/very good health-excluded-), insurance status (insured), census region (Northeast, Midwest, West, South-excluded), presence of father in the household, presence of other adults in the household, having a health limitation, residence in a metropolitan statistical area, whether the mother is the survey respondent and a constant. The p value for the F statistic is given in the bracket.

Table 3.6 Impact of Maternal Work on Number of Curative Visits

Table 3.6 Impact of Maternal Work on Number of Curative Visits

	(1)		(2)		(3)	
Mother's hours of work/week	-0.0061**	(0.0030)	-	-	-0.0064**	(0.0027)
Mother's earnings /week	0.0002**	(0.0001)	0.0002**	(0.0001)	0.0001	(0.0001)
Mother working	0.0176	(0.1054)	0.0013	(0.0919)	0.0133	(0.1050)
Mother's self-employment hours/week	-	-	-0.0019	(0.0020)	-	-
Mother's wage-employment hours/week	-	-	-0.0051**	(0.0025)	-	-
Mother's hours x child age <2	-	-	-	-	0.0035	(0.0038)
Mother's hours x child age 3-5	-	-	-	-	0.0002	(0.0039)
Mother's earnings x child age <2	-	-	-	-	0.0001	(0.0002)
Mother's earnings x child age 3-5	-	-	-	-	0.0002	(0.0002)
Other household income/week	0.00004	(3.3e-05)	0.00004	(3.3e-05)	0.00004	(3.3e-05)
Number of observations	19,390		19,390		19,390	
F statistic for test of equal income effects	1.94 [0.16]		1.24 [0.26]		0.79 [0.37]	

Notes: The Negative Binomial Regression results for the number of annual curative care visits of children. Robust standard errors are in parentheses.*** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All specifications also include mother's age and education, family size, indicator variables for child's age (less than 2, 3 to 5 and 6-12-excluded-), sex (male), race/ethnicity (Hispanic, black and white -excluded-), health status (fair or poor health, good health and excellent/very good health-excluded-), insurance status (insured), census region (Northeast, Midwest, West, South-excluded), presence of father in the household, presence of other adults in the household, having a health limitation, residence in a metropolitan statistical area, whether the mother is the survey respondent and a constant. The p value for the F statistic is given in the bracket.

Table 3.7 Net Effects of Maternal Employment

Table 3.7 Net Effects of Maternal Employment

Probability of any preventive visit	-0.0021***	(0.0007)
Probability of any curative visit	-0.0008***	(0.0003)
Number of preventive visits	-0.0061***	(0.0019)
Number of curative visits	-0.0085**	(0.0039)

Notes: Net effects of an additional hour of maternal work on children's health care utilization, calculated at the sample mean hourly earnings of working mothers ($12).

Table 4.1 Sample Statistics

Table 4.1 Sample Statistics

	Children		Wives		Husbands	
	Mean	Standard Deviation	Mean	Standard Deviation	Mean	Standard Deviation
Number of preventive visits	0.96	(1.25)	1.38	(2.92)	0.65	(1.29)
Mother's non-labor income	3,759	(9,575)	-	-	-	-
Father's non-labor income	3,100	(7,768)	-	-	-	-
Mother's education	13.20	(3.04)	-	-	-	-
Father's education	13.09	(3.12)	-	-	-	-
Non-labor income	-	-	4,315	(10,323)	3,650	(8,496)
Spouse's non-labor income	-	-	4,210	(9,105)	4,000	(9,989)
Years of education	-	-	13.30	(2.85)	13.15	(2.98)
Spouse's years of education	-	-	13.24	(2.93)	13.27	(2.90)
Fair/poor health	0.02	(0.13)	0.10	(0.30)	0.08	(0.28)
Health limitation	0.03	(0.16)	0.16	(0.37)	0.15	(0.35)
Uninsured	0.06	(0.24)	0.10	(0.31)	0.12	(0.33)
Male	0.52	(0.50)	-	-	-	-
Age	7.77	(4.75)	39.03	(9.19)	40.15	(8.67)
Family size	4.43	(1.12)	3.26	(1.23)	3.34	(1.24)
Hispanic	0.19	(0.39)	0.15	(0.35)	0.16	(0.37)
Black	0.08	(0.27)	0.07	(0.26)	0.07	(0.26)
Mother is survey respondent	0.82	(0.38)	-	-	-	-
Survey respondent	-	-	0.77	(0.42)	0.22	(0.41)
Number of observations	6,054		4,523		4,137	

Table 4.2 Determinants of Children's Preventive Care Utilization

Table 4.2 Determinants of Children's Preventive Care Utilization

Mother's non-labor income	0.0091***	(0.0019)
Father's non-labor income	0.0006	(0.0015)
Mother's education	0.0438***	(0.0031)
Father's education	0.0340***	(0.0101)
Fair/poor health	0.1595*	(0.0912)
Health limitation	0.5524***	(0.0553)
Uninsured	-0.3902***	(0.0538)
Male	0.0125	(0.0356)
Age	-0.0760***	(0.0052)
Family size	-0.0269***	(0.0089)
Hispanic	-0.0594***	(0.0168)
Black	-0.2768***	(0.0474)
Mother is survey respondent	0.1123**	(0.0497)
Constant	-0.5170**	(0.2153)
Number of observations	6,054	
Equal income effects chi-sq(1)	3.99**	[0.028]
Equal education effects chi-sq(1)	1.98	[0.322]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of children. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.3 Interaction Effects Between Parental Resources and Child Gender

Table 4.3 Interaction Effects Between Parental Resources and Child Gender

Mother's non-labor income	0.0073***	(0.0013)
Father's non-labor income	-0.0004	(0.0021)
Mother's education	0.0482***	(0.0039)
Father's education	0.0268***	(0.0011)
Mother's non-labor income*Male	-0.0006	(0.0010)
Father's non-labor income*Male	-0.0052	(0.0031)
Mother's education*Male	0.0184	(0.0137)
Father's education*Male	-0.0216	(0.0158)
Number of observations	6,054	
Equal income effects chi-sq(1)	3.78**	[0.051]
Equal education effects chi-sq(1)	2.25	[0.133]
Joint significance of interaction terms chi-sq(4)	4.18	[0.123]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of children. The model also includes child's age, family size, indicators for fair/poor health, having a health limitation, being uninsured, male, Hispanic, black, survey respondent and a constant. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.4 Determinants of Children's Preventive Care Utilization-Only children with positive parental non-labor incomes

Table 4.4 Determinants of Children's Preventive Care Utilization-Only children with positive parental non-labor incomes

Mother's non-labor income	0.0079***	(0.0015)
Father's non-labor income	-0.0002	(0.0013)
Mother's education	0.0357***	(0.0007)
Father's education	0.0363***	(0.0059)
Number of observations	2,357	
Equal income effects chi-sq(1)	4.21**	[0.040]
Equal education effects chi-sq(1)	0.02	[0.899]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of children. Sample is restricted to the children whose both parents report positive non-labor income. The model also includes child's age, family size, indicators for fair/poor health, having a health limitation, being uninsured, male, Hispanic, black, survey respondent and a constant. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.5 Determinants of Children's Preventive Care Utilization-Parental asset incomes as power measures

Table 4.5 Determinants of Children's Preventive Care Utilization-Parental asset incomes as power measures

Mother's asset income	0.0050**	(0.0021)
Father's asset income	-0.0018	(0.0012)
Mother's education	0.0439***	(0.0032)
Father's education	0.0346***	(0.0113)
Number of observations	6,054	
Equal income effects chi-sq(1)	3.21**	[0.044]
Equal education effects chi-sq(1)	0.75	[0.386]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of children. Parental asset incomes replace non-labor incomes. Asset income consists of interest, dividend, trust and rent income. The model also includes child's age, family size, indicators for fair/poor health, having a health limitation, being uninsured, male, Hispanic, black, survey respondent and a constant. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.6 Determinants of Spouses' Preventive Care Utilization

Table 4.6 Determinants of Spouses' Preventive Care Utilization

	Wife		Husband	
Non-labor income	0.0044***	(0.0012)	-0.0010	(0.0007)
Spouse's non-labor income	0.0001	(0.0003)	0.0041	(0.0027)
Years of education	0.0296***	(0.0026)	0.0047 **	(0.0023)
Spouse's years of education	0.0059**	(0.0030)	0.0257 **	(0.0107)
Fair/poor health	0.2012***	(0.0192)	0.2593***	(0.0114)
Health limitation	0.2162***	(0.0266)	0.2060***	(0.0153)
Uninsured	-0.3873***	(0.0564)	-0.7727***	(0.0320)
Age	0.0170***	(0.0013)	0.0339***	(0.0016)
Family size	-0.0276***	(0.0010)	-0.0403***	(0.0016)
Hispanic	-0.1905***	(0.0472)	-0.3374***	(0.0475)
Black	-0.1818***	(0.0356)	-0.1849***	(0.0751)
Survey respondent	0.2339***	(0.0179)	0.2547***	(0.0195)
Constant	-0.7978***	(0.0465)	-2.0548***	(0.0823)
Number of observations	4,523		4,137	
Equal income effects chi-sq(1)	5.75**	[0.019]	2.54	[0.270]
Equal education effects chi-sq(1)	4.60**	[0.020]	5.14**	[0.048]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of wives and husbands. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.7 Determinants of Spouses' Preventive Care Utilization- Only spouses with positive non-labor incomes

Table 4.7 Determinants of Spouses' Preventive Care Utilization-
Only spouses with positive non-labor incomes

	Wife		Husband	
Non-labor income	0.0056***	(0.0011)	-0.0022*	(0.0013)
Spouse's non-labor income	0.0003	(0.0020)	0.0047	(0.0036)
Years of education	0.0310***	(0.0029)	0.0094***	(0.0030)
Spouse's years of education	0.0072**	(0.0028)	0.0349**	(0.0139)
Number of observations	2,043		1,826	
Equal income effects chi-sq(1)	7.73**	[0.039]	4.00	[0.160]
Equal education effects chi-sq(1)	5.26**	[0.048]	6.25**	[0.045]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of wives and husbands. Sample is restricted to the spouses with positive non-labor incomes. The model also includes age, family size, indicators for fair/poor health, having a health limitation, being uninsured, Hispanic, black, survey respondent and a constant. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Table 4.8 Determinants of Spouses' Preventive Care Utilization-Spouses' asset incomes as power measures

Table 4.8 Determinants of Spouses' Preventive Care Utilization-Spouses' asset incomes as power measures

	Wife		Husband	
Asset income	0.0057	(0.0036)	-0.0031	(0.0025)
Spouse's asset income	-0.0024	(0.0048)	0.0043	(0.0038)
Years of education	0.0271***	(0.0033)	0.0050**	(0.0022)
Spouse's years of education	0.0060**	(0.0030)	0.0212**	(0.0092)
Number of observations	4,523		4,137	
Equal income effects chi-sq(1)	2.81	[0.233]	2.12	[0.456]
Equal education effects chi-sq(1)	6.02**	[0.044]	3.89*	[0.082]

Notes: Negative Binomial Regression results for the number of annual preventive care visits of wives and husbands. Spouses' asset incomes replace their non-labor incomes. Asset income consists of interest, dividend, trust and rent income. The model also includes age, family size, indicators for fair/poor health, having a health limitation, being uninsured, male, Hispanic, black, survey respondent and a constant. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. Numbers in brackets are the p values of the chi-square test statistics.

Appendices

APPENDIX A: CHAPTER 2

Table A.1 Determinants of Compliance Ratio-First Stage Estimates

Table A.1 Determinants of Compliance Ratio-First Stage Estimates

	Ordinary Least Squares Estimates	
Any private insurance	8.3185**	(4.0950)
Only public insurance	6.7652**	(3.1630)
Northeast	13.3247***	(4.9655)
Midwest	11.0707***	(2.9546)
West	2.0120	(3.4112)
Some history of a serious illness in year 1	7.7687**	(3.3121)
Number of reported health conditions	1.6073***	(0.5146)
Age in months	-0.9282***	(0.1627)
Male	-3.3456	(2.7682)
Hispanic	-11.5984***	(3.2864)
Black	-7.6991**	(3.7315)
Family income/10,000	0.6096**	(0.2956)
Family size	-4.4105***	(0.9807)
Mother employed	-3.2435	(2.4585)
Mother's years of education	1.1318***	(0.3921)
Mother's age	0.7992***	(0.2270)
Father not present	-0.2431	(2.9448)
MSA	1.8032**	(0.7801)
Constant	46.6462***	(13.4290)
Number of observations	1,729	
R^2	0.293	

Notes: Robust standard errors are in parenthesis. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level. All of the explanatory variables are measured in the first-year survey.

APPENDIX B: CHAPTER 3

Table B.1 Impact of Maternal Work on Probability of Any Preventive Visit

Table B.1 Impact of Maternal Work on Probability of Any Preventive Visit

Mother's hours of work/week	-0.0029***	(0.0006)
Mother's earnings/week	0.0001***	(2.0e-05)
Mother working	0.0163	(0.0223)
Other household income/week	0.00002***	(5.2e-06)
Other adults present in the household	0.0283**	(0.0121)
Mother's age	0.0019***	(0.0007)
Mother's education	0.0136***	(0.0017)
Father present	0.0457***	(0.0102)
Child age <2	0.3895***	(0.0142)
Child age 3-5	0.1455***	(0.0127)
Male	-0.0115	(0.0090)
Hispanic	-0.0065*	(0.0038)
Black	-0.0365**	(0.0164)
Fair or poor health	0.0543**	(0.0272)
Good health	0.0244**	(0.0105)
Health limitation	0.1105***	(0.0202)
Family size	-0.0085***	(0.0016)
Northeast	0.1414***	(0.0192)
Midwest	0.0568***	(0.0136)
West	-0.0131	(0.0153)
MSA	0.0272**	(0.0138)
Mother is respondent	0.0940***	(0.0112)
Insured	0.0441***	(0.0164)
Constant	-0.0569	(0.0383)
Number of observations	19,390	

Notes: The Ordinary Least Squares estimation results for the probability of at least one preventive visit during the year. Robust standard errors are in parentheses.*** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level.

Table B.2 Marginal Effects of Maternal Work Variables on Number of Preventive Visits

Table B.2 Marginal Effects of Maternal Work Variables on Number of Preventive Visits

	(1)		(2)		(3)	
Mother's hours of work/week	-0.0080***	(0.0021)	-	-	-0.0097***	(0.0029)
Mother's earnings /week	0.0002***	(4.1e-05)	0.0002***	(4.9e-05)	0.0003***	(0.0001)
Mother working	0.0515	(0.0484)	0.0367	(0.0498)	0.0505	(0.0479)
Mother's self-employment hours/week	-	-	0.0007	(0.0012)	-	-
Mother's wage-employment hours/week	-	-	-0.0076***	(0.0025)	-	-
Mother's hours x child age <2	-	-	-	-	0.0051***	(0.0015)
Mother's hours x child age 3-5	-	-	-	-	0.0020**	(0.0008)
Mother's earnings x child age <2	-	-	-	-	-0.0003***	(0.0001)
Mother's earnings x child age 3-5	-	-	-	-	-0.0001**	(4.9e-05)
Other household income/week	0.00003***	(1.1e-05)	0.00003***	(1.1e-05)	0.00003***	(1.1e-05)
Number of observations	19,390		19,390		19,390	

Notes: Marginal effects of the negative binomial coefficients presented in Table 3.4. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level.

Table B.3 Marginal Effects of Maternal Work Variables on Number of Curative Visits

Table B.3 Marginal Effects of Maternal Work Variables on Number of Curative Visits

	(1)		(2)		(3)	
Mother’s hours of work/week	-0.0121**	(0.0058)	-	-	-0.0127**	(0.0058)
Mother’s earnings income/week	0.0003**	(0.0001)	0.0002**	(0.0001)	0.0002	(0.0002)
Mother working	0.0292	(0.1739)	0.0022	(0.1530)	0.0221	(0.1736)
Mother’s self-employment hours/week	-	-	-0.0032	(0.0033)	-	-
Mother’s wage-employment hours/week	-	-	-0.0109**	(0.0055)	-	-
Mother’s hours x child age <2	-	-	-	-	0.0058	(0.0064)
Mother’s hours x child age 3-5	-	-	-	-	0.0003	(0.0065)
Mother’s earnings x child age <2	-	-	-	-	0.0002	(0.0003)
Mother’s earnings x child age 3-5	-	-	-	-	0.0003	(0.0004)
Other household income/week	0.0001	(0.0001)	0.0001	(0.0001)	0.0001	(0.0001)
Number of observations	19,390		19,390		19,390	

Notes: Marginal effects of the negative binomial coefficients presented in Table 3.6. Robust standard errors are in parentheses. *** indicates significance at the 1% confidence level, ** at the 5% confidence level and * at the 10% confidence level.

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