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Three Essays in Labor Economics

A dissertation submitted in partial satisfaction of the
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by

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Committee in charge:

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The dissertation of Patricia K. Tong is approved, and it is acceptable in quality and form for publication on microfilm and electronically:

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DEDICATION

To my advisor, Julie B. Cullen, for her enthusiasm, generosity, and never ending encouragement.

To my parents, Esther and Timothy Tong, for their love and support.

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ABSTRACT OF THE DISSERTATION

Three Essays in Labor Economics

by

Patricia K. Tong

Doctor of Philosophy in Economics

University of California, San Diego, 2010

Professor Julie B. Cullen, Chair

This dissertation consists of three separate papers studying policy relevant questions in labor economics. The goal of these studies is to understand the welfare effects from various government regulation and programs.

Chapter 1 investigates how child support income affects household resources for single mothers with at least some college education. I determine that child support income promotes single mother financial independence by reducing welfare participation, decreasing cohabitation rates, and increase labor supply.

Chapter 2 examines how a change in minimum nurse staffing regulation affects nurse employment and patient mortality in California nursing homes. My results show that regulation induced increases in nurse staffing cause patient mortality to fall by 4.6%.

Chapter 3 determines whether child support enforcement reform coming from the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 has an impact on rates of single motherhood in the United States. I find that child support enforcement reform causes the likelihood of being a single mother to increase among women with 12 or less years of education.

1 Child Support and Economic Well-Being

This study investigates the effect of child support income on household resources of single mothers with at least some college education. I exploit state level variation in child support enforcement policy to estimate the impact of child support income on different components of single mother household income. Results indicate that child support reduces welfare income, increases labor supply, and decreases the likelihood of cohabitation. Although overall household income is not significantly affected, child support income causes an increase in single mother specific resources, suggesting that single mother family well-being may be affected by a promotion in financial independence.

1.1 Introduction

This study provides a comprehensive evaluation of how child support income affects household resources of single mothers with at least some college education. Research on government transfers determines that single mothers alter their behavior in response to these additional sources of income.[1] To better understand how government transfers affect well-being, the literature tests whether transfers are crowded out by other sources of income (Gruber, 1997; Cullen and Gruber, 2000; Gruber, 2000; Engen and Gruber, 2001) and how material well-being of single mothers has changed over time (Meyer and Sullivan, 2004). This paper adds to the existing literature by examining how child support, a different type of transfer,

[1]See Moffitt (1992) and Blank (2002) for comprehensive literature reviews on welfare.

affects single mother behavior and well-being.

Unlike government transfers, child support is a government mandated private transfer of income between non-custodial and custodial parents. To qualify for child support payments, an individual must have at least one biological child aged 18 or under and must not be married to the other biological parent. Both unmarried and remarried custodial parents qualify for payments. The purpose of child support is to secure the financial well-being of children who do not live with both biological parents. Descriptive analyses demonstrate that child support income is associated with reducing poverty rates of single mother families (Nichols-Casebolt, 1986; Meyer and Hu, 1999; Bartfeld, 2000). Child support also positively correlates with scholastic outcomes including test scores, educational attainment, and grade point average (Graham et al., 1994; Knox and Bane, 1994; McLanahan et. al, 1994; Knox, 1996; Argys et al., 1998). An implicit assumption in the existing literature is that child support income translates to more financial resources, with a dollar increase in child support resulting in an additional dollar of income for the single mother. However, this is only true if child support income does not affect other sources of income. In this paper, I determine how child support income affects overall household income and the different components of single mother household income.

To identify the causal effect of child support income on single mother resources, I use state level variation in the implementation of child support reform to instrument for child support income. This study investigates outcomes for single mothers with at least some college education, which make up 40% of the single mother population. From this point forward, I refer to single mothers with at least some college education as high educated single mothers. I exclude low educated single mothers from the study because the first stage is not identified for this sample.[2] Despite their additional years of education, high educated single mothers have significantly fewer resources than the average American household, with poverty rates over 1.5 times the national average in 2007. By evaluating the effect of child support income on household resources, I provide implications on

[2]The instrumental variables estimation for single mothers with a high school degree or less suffers from a weak instrument problem.

how child support affects the circumstances in which children of high educated single mothers are raised.

Results show that a $1 increase in child support income per year causes a 25 cent reduction in welfare income, revealing that there is redistribution between public and private transfers. This redistribution suggests that the existing welfare literature may present an incomplete picture if an evaluation of child support is absent. Single mothers use this transfer of income to purchase their independence, with an additional $1000 of child support income resulting in a 4.64 percentage point reduction in the likelihood of cohabiting with an unmarried male partner. Single mothers who are induced to live independently also respond by increasing their labor supply. Household income is not significantly affected by child support income. However, child support income promotes single mother financial independence by decreasing welfare benefits, reducing cohabitation rates, and increasing single mother labor supply.

This paper will be organized as follows. Section 2 provides background on child support enforcement policy, Section 3 explains the predicted effects of child support income, Section 4 describes the data, Section 5 details the identification strategy, Section 6 discusses the empirical estimation and results, and Section 7 concludes.

1.2 Background

While the federal Office of Child Support Enforcement (CSE) states their goal as securing the "well-being of children by assuring that assistance in obtaining support...is available to children," federal intervention in child support collection originated as a way to make sure welfare participants were not receiving both welfare benefits and child support. In 1974, federal and state CSE offices were created to collect child support for custodial parents on welfare. Welfare participants were required to comply with CSE agencies in order to receive benefits, and state governments retained child support collections, causing non-custodial parents to have little incentive to make payments. The 1984 CSE amendments expanded services to non-welfare participants. In recognition of the disincentives associated with child support payments in welfare cases, the Deficit Reduction Act of 1984 required states to pass the first $50 of monthly child support payments collected on behalf of welfare participants to the custodial parent. This $50 is commonly referred to as the child support pass through.

The most recent wave of child support legislation occurred with the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996, a comprehensive welfare reform act. Under PRWORA, states were required to institute reforms to improve paternity establishment rates, impose better technology to locate parents, streamline the imposition of penalties, and create new penalties. States were also allowed to change their child support pass through policies. Non-compliance did not result in direct fines; however, states did have financial incentives to comply because federal funding of state welfare programs was directly tied with child support performance. In the empirical section, I exploit state level variation in the implementation of CSE policies created by the PRWORA to identify the causal impact of child support income on single mother resources.

1.3 Predicted Effects

In this section, I discuss the different ways that child support income could impact the three main sources of high educated single mother household income: 1) single mother non-wage income, 2) income from other household residents, and 3) single mother wage income. By investigating how child support affects non-wage income, I determine whether child support income crowds out other transfer payments. I examine the effect of child support income on living arrangements and single mother labor supply to test whether high educated single mothers adjust their behavior in response to a change in net income. Although the empirical analysis is restricted to the sample of high educated single mothers, the predicted effects of child support income apply to both high and low educated single mothers. As a result, I will discuss the predicted effects of child income for single mothers in general and refer to high educated single mothers when appropriate.

1.3.1 Single Mother Non-Wage Income

In addition to child support income, high educated single mothers have two other main sources of non-wage income, welfare benefits and other government transfers.[3] High educated single mothers receive 10% of their non-wage income through welfare benefits and 20% through other government transfers. In total, this income accounts for 25% of high educated single mother non-wage income or approximately $1420. Average child support income is $2760, which accounts for 50% non-wage income.

Child support income is predicted to decrease welfare income. As discussed, there is a direct negative relationship between welfare benefits and child support income because states retain the majority of child support income collected on behalf of welfare participants. In addition, single mothers may become ineligible for welfare if child support payments are large enough to cause income to exceed the means-tested threshold. Increasing child support payments might also change a single mother's decision to participate in welfare. An increase in expected child support payments could deter a single mother from participating if she believes that these payments will exceed welfare benefits. After 1996, welfare participation was limited to five years, causing single mothers to strategically decide when to participate (Grogger, 2002; Grogger, 2004; Mazzolari, 2007). Similarly, a single mother may decide to postpone welfare participation when she receives child support as a way to insure against a future reduction in income.

The predicted effect of child support income on other government transfers is likewise negative because a single mother is less likely to qualify for means-tested programs when she receives child support. All other transfers are predicted to be unaffected by child support income because they require some event to occur. For

[3]Other government transfers include supplemental security income, unemployment benefits, workers compensation, veteran's benefits, disability benefits, and educational assistance. Other sources of non-wage income such as alimony, investment income, and retirement income are excluded from the analysis because receipts of these types of income does not occur frequently enough to be identified by state level variation.

example, a single mother must be laid off to qualify for unemployment benefits or experience a work related injury to receive workers compensation.

1.3.2 Income from Other Household Residents

Nearly 30% of high educated single mothers share a residence with individuals in addition to her biological children. Approximately 10% of high educated single mothers cohabit with an unmarried male partner and 14% live with a relative. When high educated single mothers live with additional individuals, nearly half of household income is attributed to these other residents. Therefore, changes in living arrangements can potentially have a substantial impact on household resources.

In this section, I discuss reasons why child support income could increase or decrease the likelihood that a single mother shares a residence, causing the effect of child support income on living arrangements and other resident income to be ambiguous. I model the decision to live in a joint residence as a cooperative bargaining model, where the decision to share a residence is a utility maximizing decision for the single mother and other potential residents. This set up is inspired by Manser and Brown (1980) and McElroy and Horney (1981) who were the first to model family and marriage decisions with individual level utility functions. Common preference models use household utility which suggests that family consumption should be the same regardless of who controls spending (Samuelson, 1956; Becker, 1974; Becker 1991). Empirical research rejects this assertion and generally finds that mothers spend more on children than fathers (Thomas, 1990; Lundberg et al., 1997; Browning and Chiappori, 1998; Phipps and Burton, 1998). A single mother and other household residents may have different preferences for consumption, making the cooperative bargaining framework more appropriate. Moreover, using individual specific utility allows a comparison of utility under different living arrangements, which is not possible with a common preference model.

Individuals maximize utility

$$U_i = U(\bar{Y}, 1 - H_i)$$

subject to a budget constraint

$$\bar{Y} = \alpha_i Y_i + \sum_{i \neq j} \alpha_j Y_j = \vec{p}_i \vec{x}_i$$

$\bar{Y}$ represents total household income and $1 - H_i$ is hours of leisure. Household income equals a weighted sum of income from each member of the household and $\vec{p}_i \vec{x}_i$ are total expenditures. α is a resident specific number between 0 and 1 and discounts income when a residence is shared. Under a joint residence, a single mother derives a positive utility from the portion of resident j's income shared with her. Similarly, the single mother shares part of her income with other residents and is left with α_i of her own income. When single mother i lives independently, meaning she only lives with her children, $\alpha_{i \neq j}$ equals zero and α_i equals 1. A joint residence will occur if two people derive a greater utility from living together than from living apart.

There are several reasons why child support income could promote the likelihood that a single mother lives with additional people. Rises in child support income will increase resident j's utility because utility is increasing in single mother income. Furthermore, the receipt of child support might serve as a signal to potential residents that they will not have to spend their own resources to pay for child expenses because the children are already funded by their father.

In contrast, child support income may also decrease the likelihood that a single mother shares a residence. According to Becker (1973), women and men decide to marry because there are gains to specialization. These gains in specialization are based on relative wage. The partner with the higher relative wage, typically the man, specializes in market work while the other partner specializes in household production. Becker's model predicts that a woman's likelihood of marriage decreases as her relative wage increases. In accordance, studies hypothesize that an increase in a woman's economic independence will decrease the likelihood of marriage and increase divorce, otherwise known as the "independence effect" (Oppenheimer, 1997; Sayer and Bianchi, 2000). Analogously, when a single mother's economic resources increase because of child support income, she could become less likely to live with others because there are fewer gains to specialization.

Child support income might also decrease the likelihood that a single mother shares a residence by increasing the probability that she can afford to live independently. A benefit to sharing a residence is the ability to pool resources to pay for large expenses like rent. Expenses are generally lower when a single mother shares a residence than when she lives independently. When a single mother receives child support income, she becomes more likely to be able to purchase her independence.

Child support income might also change income of other residents without affecting living arrangements. For instance, higher child support payments may have a negative income effect on other residents and cause them to decrease labor supply.

This discussion, while not exhaustive, demonstrates that there are multiple ways that child support can increase and decrease income from other individuals in the household. As a result, determining how child support income affects these resources is left as an empirical question.

1.3.3 Single Mother Wage Income

Although traditional economic theory suggests that a positive non-wage income shock will decrease labor supply and wage income, this may not necessarily hold true for single mothers. Single mothers must have arrangements for child care if they decide to work during non-school hours. If child support income is used to purchase child care or other work related expenses, then it could have a positive effect on labor supply and wage income. Furthermore, studies demonstrate that increasing child support payments promotes visitation of non-custodial fathers (Weiss and Willis, 1985; Graham and Beller, 2002; Argys and Peters, 2003; Nepomnyaschy, 2007). Weiss and Willis (1985) develop a theoretical model in which children are collective goods and fathers increase time with their children in order to monitor how child support income is spent. During visitation, fathers spend additional resources on their children by providing housing on overnight visits and paying for meals. If child support promotes visitation, then this source of secondary financing might allow the single mother to increase labor supply because she has more time to work.

If increases in child support income induce single mothers to live independently, then labor supply might rise because a single mother now faces a higher marginal wage rate. Under a shared residence, a single mother loses $1 - \alpha_i$ of her wage income to other residents, but retains all of her earned income when living independently. Single mothers may also view independence as a good itself and derive a positive utility from living independently. Unlike a continuous good such as food, independence is a discrete good that can only be purchased with a sufficient amount of income. The combination of child support income *and* increases in wage income could allow a single mother to be able to afford independence.

If child support income causes a single mother to move from living independently to sharing a residence, then there are still scenarios where she could increase labor supply. With higher child support payments, a single mother might be able to increase her labor supply by moving in with her parents and compensating them for child care. This is plausible if a single mother has a greater earnings capacity than her parents. As with living arrangements, the effect of child support income on labor supply and wage income is ambiguous and left as an empirical question.

1.4 Data

In this study, I use March Current Population Survey (CPS) data for years 1992-2004. I limit the sample to single mothers with at least some college education. A single mother is defined as a divorced, separated, or never married woman with at least one biological child aged 18 or under who lives in the same household. I limit the sample to high educated single mothers who have children aged 5-18.[4] Research demonstrates that a significant fraction of couples who have out-of-wedlock births marry within 5 years, causing a high level of attrition out of single motherhood (Carlson et al., 2004, Harknett and McLanahan, 2004). This

[4]To ensure that results are not driven by outliers, I also drop observations in the top and bottom 1% of the household income distribution. In addition, I exclude 14 observations reporting more than $50,000 in child support income and exclude step families from the study because CPS data do not consistently identify step children over the time period of interest.

also applies to divorced single mothers because births may have occurred after the marital dissolution. Previous literature provides empirical evidence that couples are less likely to divorce when they have pre-school aged children (Cherlin, 1977; Waite and Lillard, 1991), implying that divorced mothers with children aged 0-4 are more likely to have experienced a non-marital birth than divorced mothers with children aged 5-18. Due to a potentially high level of attrition out of single motherhood among women with recent births, I drop high educated single mothers with children aged 4 and under.[5]

Table 1.1 reports summary statistics for high educated single mothers by child support receipt. All monetary variables are deflated to 2007 dollars and reported in annual terms. 46% of high educated single mothers receive child support and the unconditional average of child support income is $2760 per year. 1 in 4 high educated single mothers have a college degree. Summary statistics demonstrate that there are differences across observable characteristics by child support receipt. In particular, non-recipients are more likely to live in a city, be Hispanic or Black, and participate in welfare than high educated single mothers who receive child support.

1.5 Identification Strategy

To account for endogeneity in child support income, I exploit state level variation in the timing in which states institute child support reform created by the PRWORA. I use Office of Child Support Enforcement State Plans to obtain CSE law data. Although the provisions of the PRWORA apply to each state, there is variation in when these laws are passed. There are three main waves in which states implement child support reform occurring in 1997, 1998, and 1999. In Table 1.7, I categorize states as 1997 movers, 1998 movers, and 1999 movers based on the year when a state passes the most laws. There are 8 1997 movers, 33 1998

[5]Although previous studies define pre-school aged children as children aged 0 to 5, I use a more conservative age range of 0-4 because the minimum age requirement to attend kindergarten in the U.S. is 5.

movers, and 9 1999 movers.[6] States generally pass a majority of laws, over 80%, when they move.

To measure CSE reform, I construct a state-year variable called First Pass, which equals one when any laws are passed in a state and zero otherwise.[7] To provide evidence that First Pass is a valid instrument, I collapse pre-reform data for years 1994-1995 to the state level and use a multinomial logistic estimation to determine if pre-reform state level characteristics affect the likelihood that a state is a 1997, 1998, or 1999 mover. Results in Table 1.8 reveal that states with a higher fraction of women and a lower fraction of Republicans in the state legislature are more likely to be a 1997 mover relative to a 1999 mover.[8] Similarly, states with a higher fraction of women in the state legislature are more likely to be a 1998 mover relative to a 1999 mover. Having a higher fraction of women in a state legislature could cause states to implement reform sooner because women might be more sympathetic to the single mother population as possible mothers themselves. In general, Democrats are more likely to support increases in expenditures on social welfare programs than Republicans, which explains why states with a higher fraction of Democrats in the state legislature implement CSE reform sooner. Pre-reform state level child support outcomes, as measured by the fraction of child support cases with collections, do not significantly affect when states pass CSE laws. Furthermore, pre-reform state level female demographic variables have a jointly insignificant effect on state mover status. Therefore, I do not find the timing of CSE reform to be correlated with these potentially confounding factors. Instead, the implementation of CSE reform appears to be a result of pre-reform state level

[6] Indiana is omitted from this categorization because it does not institute CSE reform until 2001.

[7] I use a restricted set of laws to calculate the First Pass variable. Although I have data on 16 CSE laws, these measures are calculated using a subset of 11 laws. These 11 laws are policies in which the majority, if not all, are passed when a state begins implementing reform. The remaining 5 laws are passed sporadically across and within each state. Including these additional 5 laws in the first stage regression does not change the coefficient estimate on the First Pass variable. Additionally, redefining the First Pass variable to include all 16 laws yields quantitatively similar results, but a smaller first stage F-statistic. A description of the 11 laws included in the analysis can be found in the Appendix.

[8] Data on the fraction of women and Republicans in state legislatures were obtained from replication datasets for Berry et al. (2007) posted on http://www.uky.edu/∼rford/replicationdata.html.

political climate, which I assert is independent of child support collection.

1.6 Empirical Results

To analyze how child support income affects high educated single mother household resources, I use an instrumental variables estimation. The equation of interest is:

$$\begin{aligned} H_{ist} = & A_s + B_t + A_s t + \beta \text{Child Support}_{ist} \\ & + \Gamma X_{ist} + \pi S_{st} + \epsilon_{ist} \end{aligned} \tag{1.6.1}$$

where H_{ist} is an outcome for single mother i in state s and year t. X_{ist} is a vector of individual level characteristics including age, age squared, race, city residence status, age of youngest child, number of children aged 18 and under, and number of children older than 18. S_{st} is a vector of state time-varying characteristics including the unemployment rate, average female hourly wage, average male hourly wage, the maximum child tax credit, the maximum Earned Income Tax Credit (EITC), and the natural logarithm of the maximum annual welfare benefit for a three person household.[9] S_{st} also includes dummy variables indicating when Temporary Assistance to Needy Families[10] (TANF) reform occurs, when TANF time limits are imposed, and whether the state has a child support pass through.[11] I include state fixed effects, A_s, to account for time-invariant state level characteristics and year fixed effects, B_t, to account for time-varying characteristics shared across states. $A_s t$ are state fixed trends which account for state specific variables trending linearly during this time period.

[9] Unemployment rate was obtained from the Bureau of Labor Statistics Archived Regional and State Unemployment data. Average female and male hourly wage are calculated using March CPS Data. The maximum child tax credit is the maximum child tax credit that can be claimed per child. The maximum EITC is the maximum state and federal EITC calculated based on number of children. The maximum annual welfare benefit was obtained from the Urban Institute's Welfare Rules Database and the Department of Health and Human Services.

[10] The PRWORA of 1996 changed welfare programs and renamed the welfare program Temporary Assistance to Needy Families.

[11] TANF implementation data was obtained from Bitler et al. (2006). TANF time limit and child support pass through data was extracted from the Urban Institute Welfare Rules Database.

Using Ordinary Least Squares (OLS) to estimate the effect of child support income on single mother outcomes is problematic because there are unobservable characteristics correlated with child support income. Summary statistics by child support receipt reveal that there are differences across observable characteristics, suggesting that there might also be differences across unobservable characteristics. For instance, women who have children with higher income men are more likely to receive child support because these men are more likely to work in the formal labor market and are, therefore, easier to locate than lower income men. In addition, these women are more likely to be high income themselves due to assortative mating, which implies that women and men form unions with individuals who are similar to themselves. Because characteristics of the biological father are unobserved, OLS estimates a non-causal relationship between child support and single mother outcomes. To establish a causal relationship, I use the First Pass variable to instrument for child support income and estimate how predicted values of child support income affect high educated single mother outcomes.

1.6.1 First Stage Results

The first stage estimates the effect of CSE reform on child support income for high educated single mothers. Results reveal that CSE reform causes child support income to increase by nearly $800, which is statistically significant at the 1% level.[12,13] This result supports previous research which finds that CSE policy improves child support outcomes (Beller and Graham, 1993; Garfinkel and Robins, 1994; Freeman and Waldfogel, 2001; Sorensen and Hill, 2004). Summary statistics show that the average amount of child support income is $2760 per year, indicating that CSE reform causes a 30% increase in child support income. The partial F-statistic on First Pass is 10.75 which passes Staiger and Stock's (1997) rule of

[12] The coefficient estimate on First Pass is 792.719 and has a standard error of 241.81.

[13] To understand how the addition of state fixed effects, year fixed effects, and state fixed trends affects the coefficient estimates, I systematically added these measures to a baseline linear regression model. The First Pass variable is insignificant in the baseline model. The addition of state fixed effects causes the First Pass variable to have a significant and positive effect on child support income. Adding year fixed effects and state fixed trends to the state fixed effects model causes the coefficient estimate on First Pass to increase by about 60% and the standard errors to decrease.

thumb.

1.6.2 Second Stage Results

To analyze how child support income affects high educated single mother household resources, I instrument for child support income using the First Pass variable. I report OLS estimates along with the Instrumental Variables (IV) estimates for illustrative purposes. Recall that OLS estimates give a non-causal relationship between child support income and high educated single mother outcomes because there are unobservable characteristics correlated with child support income and the outcome variables. A secondary issue when comparing OLS to IV estimates is that IV estimates are identified for individuals who experience CSE induced increases in child support income. Even if OLS estimates were causal, they would identify the effect of child support income on outcomes for the average high educated single mother, which might not be the same as the sample affected by CSE reform.

Single Mother Non-Wage Income

To determine the net effect of child support income on single mother non-wage income, I investigate how child support income affects welfare benefits and other government transfers. Estimates in Table 1.3 demonstrate that increases in child support income induced by CSE reform crowd out welfare benefits, with each dollar of child support income yielding a 25 cent reduction in welfare benefits. Therefore, an $800 rise in child support causes welfare income to fall by $200. This provides evidence of redistribution between private and public transfers, suggesting that previous research examining the effect of welfare on single mother outcomes might be incomplete if an analysis of child support income is omitted. Prior to the PRWORA, approximately 14% of high educated single mothers participated in welfare. After reform occurred, only 4% collected welfare benefits. Conditional on participating in welfare, recipients received approximately $5000 per year in welfare benefits before the PRWORA. If the $200 reduction in benefits is caused entirely by a decrease in participation, then this implies that CSE induced changes in child support income cause welfare participation to fall by 4 percentage points. In practice, increases in child support income will also cause welfare benefits to reduce among participants, meaning that a 4 percentage point decrease is an upper bound for the impact of CSE induced changes in child support on welfare participation. Other government transfers are not significantly affected by child support income. Therefore, there is a positive net income effect on high educated single mother non-wage income from increases in child support income, with each dollar of child support yielding 75 cents of non-wage income. Given that CSE reform causes an $800 increase in child support income, single mothers have an additional $600 to spend after accounting for the reduction in welfare income.

Income from Other Household Residents

Results in Table 1.4 and 1.5 report the effects of child support income on the likelihood of living with different individuals and on their respective income. While child support income does not have a significant effect on the probability of living with a relative or income from relatives, child support income does have a statistically significant and negative effect on the likelihood of cohabiting with an unmarried male partner and income from an unmarried male partner. Thus, single mothers who are leaving cohabiting households are being induced to live independently.

CSE caused child support income to increase by an average of $800, meaning that CSE induced changes in child support income decrease the likelihood that a high educated single mother cohabits with an unmarried male partner by 3.7 percentage points, a 40% reduction from the mean. These results suggest that high educated single mothers who receive more child support income use the extra income to purchase their independence. A similar result is found in a study by Bitler et al. (2006) which determines that state welfare waivers decrease the likelihood of cohabitation, but only for low educated black mothers. The literature investigating how welfare affects single mother independence as measured by the incidence of marriage and female headship also finds positive effects (Ellwood and Bane, 1985; Moffitt, 1990; Schultz, 1994); however, a more recent paper by Hoynes

(1997) determines that there is no significant impact on female headship.

An \$800 increase in child support income causes almost a \$2000 decrease in male partner income. The conditional average of male partner income is \$33,830, meaning that a \$2000 decrease is equivalent to a 6% decline from the conditional average. This estimate is conservative because single mothers who are on the margin of leaving a cohabiting relationship are likely in less financially advantageous partnerships. Average income of male partners who have income below the conditional mean is \$17,500. If high educated single mothers who are induced to live independently are dissolving relationships with men with an average income of \$17,500, then a \$2000 reduction in male partner income is equivalent to a 11.4% reduction.

I hypothesize that this reduction in male partner income is explained by the decrease in the probability of cohabiting and not by changes in male partner income. The 95% confidence interval of the coefficient estimate for child support income demonstrates that a \$1000 increase in child support income causes unmarried male partner income to decrease anywhere from \$530 to \$4430. Furthermore, a \$1000 increase in child support income causes the likelihood that a single mother cohabits with an unmarried male partner to decrease by 4.64 percentage points. If high educated single mothers are ending relationships with men who have an average income of \$33,830, then a 4.64 percentage point reduction in the likelihood of cohabiting with a male partner would translate to a \$1570 decrease in male partner income, falling well within the 95% confidence interval. Similarly, if high educated single mothers are ending relationships with men who have an average income of \$17,500, then a 4.64 percentage point reduction in the likelihood of cohabiting is equivalent to an \$812 decrease in male partner income, also within the 95% confidence interval. This accounting exercise provides evidence that the decrease in the probability of cohabiting likely explains the decrease in male partner income.

Single Mother Wage Income

Results in Table 1.6 show that increases in child support income induced by child support reform cause an increase in high educated single mother wage income. While the extensive margin of labor supply is not significantly affected, child support income causes a significant increase in the intensive margin of labor supply. In particular, an additional \$800 in child support income causes weeks worked per year to increase by an average of 1.6. Unreported results also show that the addition of child support income increases the probability that a high educated single mother has employer provided health insurance and pensions, indicating that high educated single mothers are choosing to work at better jobs or are qualifying for employer benefits through an increase in weeks worked.

Examining the coefficient estimates on other covariates shows that the EITC promotes the likelihood a single mother is employed and the child tax credit increases wage income. Although TANF reform has been found to promote labor supply of single mothers in previous literature (Meyer and Rosenbaum, 2001; Grogger, 2003), I do not find the implementation of TANF or TANF time limits to have a significant effect on either the extensive or intensive margin of labor supply. This discrepancy may be explained by the fact that most high educated single mothers do not meet the means-tested threshold, and are therefore, not as sensitive to changes in welfare policy. Furthermore, high educated single mothers who exit welfare because of increased child support payments are perhaps individuals who are already working and capable of supporting their families without welfare.

A \$1 increase in child support income increases high educated single mother wage income by \$2.13. This estimate must be interpreted with caution because it is imprecisely measured and only significant at the 10% level. This coefficient estimate implies that an \$800 increase in child support income causes a \$1700 increase in wage income. However, the 95% confidence interval of the coefficient estimate demonstrates that the change in wage income is anywhere between -\$16 and \$3400. Although the effect on wage income is imprecisely estimated, these results provide evidence that child support income causes high educated single mothers to increase labor supply, and subsequently, earn more income.

Aggregate Income

Unreported reports the estimated effect of child support income on aggregate household income measures, including total household income, household income equivalent, and total single mother income.[14] Household income equivalent is calculated based on the Organization for Economic Co-operation and Development (OECD) Household Equivalent which weights household income by the number of people in the household. The first adult is given a weight of 1, each additional adult is given a weight of 0.7, and each child is given a weight of 0.5. Increases in child support income do not significantly affect total household income. Although the estimate of the impact of child support on household income equivalent is too imprecise to make inferences, there is a large statistically significant effect of child support income on single mother income. Each dollar of child support income causes a $2.63 increase in single mother income, indicating that single mother specific resources rise as a result of child support income.

1.6.3 Robustness Checks

As a robustness check, I conduct the empirical estimation for different sample splits by age of the youngest child to determine whether the results depend on including high educated single mothers with young children. The main analysis includes high educated single mothers with children aged 5 to 18. Expanding and contracting this sample to include high educated single mothers with children aged 3 to 18 and 7 to 18 does not change the results. This implies that high educated single mothers with young children, who are the most susceptible to attrition through marriage, do not drive the results discussed in the previous sections.

As a placebo test for the first stage, I segment the data into pre and post CSE reform time periods, 1992-1996 and 2000-2004, and estimate the effect of the leads and lags of the First Pass variable on child support income in both time periods. By using the leads and lags, I institute a fake CSE policy change in the pre and post period. To pass the placebo test, there should be no effect from the

[14]Total single mother income includes child support income.

First Pass variable on child support income in either period. Unreported results confirm that there is no significant affect from the leads and lags of the First Pass variable.

1.7 Conclusion

In this paper, I examine how child support income affects monetary resources of single mothers with at least some college education. I determine that child support income crowds out welfare benefits. This redistribution between private and public transfers suggests that child support income and welfare transfers should be studied together. High educated single mothers use this additional income to purchase their independence, with CSE reform induced child support payments causing a 3.7 percentage point reduction in the probability of cohabiting with an unmarried male partner. High educated single mothers who use child support income to purchase independence also increase their labor supply and are more likely to have employer provided benefits.

Child support has no significant effect on total household income. Instead, child support income increases single mother financial independence through changes in welfare dependency, living arrangements, and labor supply. While literature examining outcomes of children living with cohabiting step parents is sparse, some studies find that these children experience more behavioral and psychological problems (Amato and Keith, 1991), and have a greater likelihood of being expelled from school than children in single parent households (Nelson et al., 2001). This suggests that reducing cohabitation rates through child support could improve child outcomes. In addition, CSE reform increases payments for high educated single mothers who are on the margin of living independently. If these single mothers are able to leave less desirable relationships and obtain jobs that provide more benefits, then increases in child support payments may improve child well-being through these different channels. Moreover, increasing a single mother's attachment to the labor force might also improve child outcomes if this creates greater employment and income stability for the household. To understand whether these changes improve child well-being, future research should investigate the relationship between single mother financial independence and child outcomes directly.

Table 1.1: Summary Statistics of Demographic Characteristics For High Educated Single Mothers

	All		Yes Child Support		No Child Support	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Child support income	2760.816	(4959.739)	6036.1	(5832.095)		
Receive child support	0.457	(0.498)				
Never married	0.252	(0.434)	0.149	(0.356)	0.339	(0.474)
On Welfare	0.095	(0.293)	0.057	(0.232)	0.127	(0.333)
Live in a central city	0.303	(0.459)	0.237	(0.425)	0.358	(0.479)
Number kids > 18	0.08	(0.302)	0.064	(0.264)	0.092	(0.33)
Age	36.769	(5.258)	37.119	(5.113)	36.473	(5.36)
Age of youngest child	10.243	(3.72)	10.075	(3.58)	10.385	(3.827)
Number of kids age ≤ 18	1.614	(0.774)	1.669	(0.744)	1.567	(0.796)
Some college	0.747	(0.435)	0.727	(0.445)	0.764	(0.425)
College graduate	0.253	(0.435)	0.273	(0.445)	0.236	(0.425)
Hispanic	0.076	(0.265)	0.061	(0.239)	0.088	(0.284)
Black	0.294	(0.455)	0.176	(0.38)	0.393	(0.488)
Observations	14422		6759		7663	

Notes: Summary statistics are constructed using March Current Population Survey data for years 1992-2004. Means are weighted by person weights. Monetary are data reported in 2007 dollars.

Table 1.2: Summary Statistics of Outcome Variables For High Educated Single Mothers

	All		Yes Child Support		No Child Support	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Household Income	44332.722	(29464.193)	47776.186	(28774.638)	41430.143	(29726.71)
Household Income Equivalent	20629.283	(13404.53)	22658.345	(13226.042)	18918.938	(13315.944)
Mom's Income	34440.314	(22700.389)	39206.508	(22726.999)	30422.775	(21887.458)
Mom's Wage Income	28827.925	(21588.083)	30342.869	(20646.316)	27550.942	(22272.532)
Relative Income	3628.285	(13189.075)	2559.758	(11232.178)	4528.971	(14576.78)
Male Partner Income	3196.783	(12029.086)	3363.087	(12505.933)	3056.6	(11610.893)
Live with Relatives (d)	0.142	(0.349)	0.093	(0.291)	0.182	(0.386)
Live with Male Partner (d)	0.094	(0.293)	0.091	(0.288)	0.097	(0.296)
Employed (d)	0.847	(0.36)	0.88	(0.325)	0.819	(0.385)
Employer Health Insurance (d)	0.604	(0.489)	0.652	(0.476)	0.563	(0.496)
Employer Pension (d)	0.449	(0.497)	0.5	(0.5)	0.407	(0.491)
Weeks Worked	48.374	(9.966)	48.624	(9.447)	48.147	(10.409)
Welfare Income	422.036	(1684.396)	223.285	(1177.363)	589.568	(1999.813)
Other Government Transfers	994.146	(3732.977)	984.355	(3471.856)	1002.398	(3939.865)
Observations	14422		6759		7663	

Notes: Summary statistics are constructed using March Current Population Survey data for years 1992-2004. Means are weighted by person weights. Monetary data are reported in 2007 dollars. Dummy variables are denoted by (d).

Table 1.3: Effect of Child Support Income on Single Mother Transfers

Variables	Welfare Income	Other Government Transfers
OLS		
Child support income	-0.038***	-0.007
	(0.005)	(0.005)
IV		
Child support income	-0.252**	-0.325
	(0.105)	(0.258)
TANF Implemented	-128.928	258.758
	(131.470)	(384.666)
LN(Welfare benefit)	-135.944	-489.591
	(389.307)	(904.151)
TANF Time Limit	203.563**	228.191
	(95.076)	(188.319)
Child Support Pass Through	90.782	349.544
	(110.845)	(311.403)
Child Tax Credit (in $100s)	-27.696***	-13.187
	(8.448)	(15.601)
EITC Credit (in $100s)	8.487	12.080
	(6.899)	(15.729)
Age	41.364	-87.968
	(39.779)	(64.444)
Age of youngest child	-48.175***	-43.904
	(15.130)	(27.885)
# of Children Age $\leq$ 18	457.241***	300.103**
	(75.217)	(151.520)
Hispanic	-95.404	-625.989
	(181.240)	(386.293)
Black, Non-Hispanic	-297.386	-717.071
	(197.978)	(556.422)
Observations	14422	14422

Notes: State clustered robust standard errors reported in parentheses. * significant at the 10% level, ** significant at the 5% level, and *** significant at the 1% level. State fixed effects, years effects, and state fixed trends included. Coefficient estimates for selected single mother demographics and state characteristics not included for brevity.

Table 1.4: Effect of Child Support Income on Single Mother Living Arrangements

Variables	Live With Relative(s)	Live With Male Partner
OLS		
Child support income	-0.0040***	-0.0023***
	(0.0006)	(0.0004)
IV		
Child support income	-0.0199	-0.0464**
	(0.0189)	(0.0221)
TANF Implemented	0.0233	0.0017
	(0.0219)	(0.0287)
LN(TANF benefit)	0.0525	0.0348
	(0.0618)	(0.0670)
TANF time limit exists	0.0042	0.0337
	(0.0131)	(0.0223)
Child Support Pass Through	-0.0127	0.0071
	(0.0187)	(0.0207)
Child Tax Credit (in $100s)	-0.0003	-0.0024*
	(0.0015)	(0.0014)
EITC Credit (in $100s)	-0.0009	0.0026*
	(0.0012)	(0.0013)
Age	-0.0093	0.0084
	(0.0092)	(0.0082)
Age of youngest child	-0.0004	-0.0021
	(0.0023)	(0.0024)
# of Children Age $\leq$ 18	-0.0079	0.0206
	(0.0127)	(0.0140)
Hispanic	0.0362	-0.0801**
	(0.0268)	(0.0393)
Black, Non-Hispanic	0.0375	-0.1530***
	(0.0362)	(0.0462)
Observations	14422	14422

Notes: State clustered robust standard errors reported in parentheses. * significant at the 10% level, ** significant at the 5% level, and *** significant at the 1% level. State fixed effects, years effects, and state fixed trends included. Coefficient estimates for selected single mother demographics and state characteristics not included for brevity. Child support income in $1000s.

Table 1.5: Effect of Child Support Income on Income From Relatives and Male Partners

Variables	Income From Relative(s)	Income From Male Partner
OLS		
Child support income	-0.095***	-0.045**
	(0.020)	(0.021)
IV		
Child support income	-0.115	-2.479**
	(0.754)	(0.994)
TANF Implemented	1,489.779*	127.744
	(826.189)	(1,265.858)
LN(TANF benefit)	985.089	-1,149.137
	(1,925.436)	(2,805.820)
TANF time limit exists	78.591	959.391
	(580.941)	(1,148.216)
Child Support Pass Through	-227.770	554.729
	(701.791)	(971.059)
Child Tax Credit (in $100s)	73.578*	-77.569
	(41.882)	(63.005)
EITC Credit (in $100s)	-69.150	144.700**
	(47.407)	(58.709)
Age	-1,189.120***	551.363
	(399.838)	(384.532)
Age of youngest child	-14.343	-180.123
	(83.208)	(112.515)
# of Children Age $\leq$ 18	-1,050.609***	859.044
	(389.230)	(663.732)
Hispanic	252.739	-3,707.575**
	(980.575)	(1,690.791)
Black, Non-Hispanic	526.410	-7,279.733***
	(1,557.435)	(2,024.309)
Observations	14422	14422

Notes: State clustered robust standard errors reported in parentheses. * significant at the 10% level, ** significant at the 5% level, and *** significant at the 1% level. State fixed effects, years effects, and state fixed trends included. Coefficient estimates for selected single mother demographics and state characteristics not included for brevity.

Table 1.6: Effect of Child Support Income on Single Mother Labor Supply

Variables	Wage Income	Employed	Weeks Worked
OLS			
Child support income	0.129*	0.0034_a^{***}	-0.000
	(0.066)	(0.0010)	(0.000)
IV			
Child support income	2.125*	-0.0043_a	0.002**
	(1.094)	(0.0276)	(0.001)
TANF Implemented	-365.952	0.0157	-0.513
	(1,477.150)	(0.0259)	(1.108)
LN(Welfare benefit)	906.496	0.0100	-0.251
	(6,182.534)	(0.0655)	(2.993)
TANF Time Limit	837.890	0.0133	0.433
	(867.828)	(0.0158)	(0.766)
Child Support Pass Through	240.088	0.0046	-0.821
	(1,296.574)	(0.0169)	(0.919)
Child Tax Credit (in $100s)	191.955**	0.0012	0.045
	(91.260)	(0.0016)	(0.056)
EITC Credit (in $100s)	-2.780	0.0033**	-0.073
	(73.554)	(0.0015)	(0.054)
Age	1,419.241***	0.0063	0.840***
	(446.154)	(0.0093)	(0.249)
Age of youngest child	297.794**	0.0020	0.276**
	(131.112)	(0.0035)	(0.127)
# of Children Age $\leq$ 18	-4,314.536***	-0.0469***	-1.948***
	(645.091)	(0.0169)	(0.612)
Hispanic	-477.574	-0.0385	2.690*
	(1,551.716)	(0.0419)	(1.379)
Black, Non-Hispanic	903.523	-0.0460	3.600*
	(2,231.592)	(0.0578)	(1.954)
Observations	14422	14422	12132
First Stage F	10.75	10.75	9.117

Notes: State clustered robust standard errors reported in parentheses. * significant at the 10% level, ** significant at the 5% level, and *** significant at the 1% level. State fixed effects, years effects, and state fixed trends included. Coefficient estimates for selected single mother demographics and state characteristics not included for brevity. Child support income in $1000s for estimates with subscript *a*.

Table 1.7: List of States by Mover Status

1997 Movers	1998 Movers		1999 Movers
Arizona	Alabama	Nebraska	Alaska
Arkansas	California	Nevada	Colorado
Connecticut	District of Columbia	New Hampshire	Delaware
Florida	Georgia	New Jersey	Kentucky
New Mexico	Hawaii	New York	Minnesota
Texas	Idaho	North Carolina	North Dakota
Vermont	Illinois	Ohio	South Dakota
West Virginia	Iowa	Oklahoma	Tennessee
	Kansas	Oregon	Wyoming
	Louisiana	Pennsylvania	
	Maine	Rhode Island	
	Maryland	South Carolina	
	Massachusetts	Utah	
	Michigan	Virginia	
	Mississippi	Washington	
	Missouri	Wisconsin	
	Montana		

Notes: States classified as a 1997 Mover if majority of child support enforcement laws are passed in 1997. 1998 and 1999 Movers are defined analogously. Indiana is omitted from this categorization since it does not institute child support enforcement reform until 2001.

Table 1.8: Effect of State Pre-Reform Characteristics on State Mover Status

Variables	1997 Mover	1998 Mover
(1) % Women in State Legislature	0.311***	0.151**
	(0.106)	(0.067)
(2) % Republicans in State Legislature	-0.121*	-0.038
	(0.063)	(0.052)
(3) % Child Support Cases With Collections	-6.183	9.958
	(10.055)	(7.207)
(4) % Women With a H.S. Degree or Less	23.817	8.493
	(20.627)	(14.111)
(5) % Women With Children	4.640	-13.585
	(31.048)	(21.183)
(6) % Women who are Single Mothers	-64.749**	-6.097
	(32.811)	(28.578)
(7) State Unemployment Rate	1.501**	1.029
	(0.715)	(0.640)
(8) State Female Hourly Wage	-0.257	-0.890
	(0.760)	(0.667)
(9) State Male Hourly Wage	-0.545	0.165
	(0.437)	(0.124)
Constant	-3.032	3.910
	(20.184)	(16.675)
Log Likelihood	-30.16772	
Pseudo R-squared	0.323	
Observations	49	
Joint Chi-Squared Statistic and P-Value:		
For (1) and (2)	10.53, 0.032	
For (4), (5), and (6)	8.99, 0.174	
For (7), (8), and (9)	11.49, 0.074	

Notes: Estimated using a multinomial logistic regression with 1999 mover as the base outcome. Raw coefficients are reported. State clustered, robust standard errors reported in parentheses. * significant at the 10% level, ** significant at the 5% level, and *** significant at the 1% level. Data for 1994-1995 collapsed into state cells.

1.8 Appendix

1.8.1 Description of CSE Laws

Some of the following provisions are amendments to previous legislation. See Title III of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 or Turetsky (1997) for more details. Title IV-D cases are cases in which the child support order is legally established through a government child support enforcement agency. Non Title IV-D cases are cases in which the child support order is established through a private lawyer or non Title IV-D agencies.

- Credit Bureaus: States must make periodic reports to consumer reporting agencies, which include the names and amounts owed by non-custodial parents with overdue payments. PRWORA (Section 367).

- Full Faith Paternity: States must acknowledge paternity establishments occurring in other states on full faith and credit. PRWORA (Section 331).

- Records: IV-D agencies must have laws that ensure that State and Federal child support agencies have access to any records used by the State for locating individuals for motor vehicle and law enforcement purposes. PRWORA (Section 325 D).

- Financial Match: States must pass laws that define how financial institutions will supply account information of non-custodial parents with child support arrears to IV-D agencies so that liens may be imposed. Matches must occur on a quarterly basis. PRWORA (Section 372).

- Income Withholding: Automatic deductions are made from wages or income to pay past-due child support. PRWORA (Section 314).

- Liens: Liens automatically arise against real and personal property (including bank accounts) for the amount of child support overdue by a non-custodial parent. PRWORA (Section 368).

- Paternity: This law expands availability of voluntary paternity acknowledgment services, allows voluntary paternity to become legal through administrative processes when paternity is uncontested, and makes the state birth record agency the repository for all paternity records. PRWORA (Sections 331-333).

- Review Orders: States must review child support orders 1) every 3 years, 2) if requested by either parent and there is a significant change in circumstances, or 3) if requested by the state agency. PRWORA (Section 351).

- Collection Social Security Numbers: States are required to collect social security numbers of any individuals who apply to get a driver's license. PRWORA (Section 317).

- Work Requirements: This law only applies to cases in which a child receives TANF. The purpose of this law is to ensure that non-custodial parents with child support arrears either work or have a plan for payment which includes appropriate work activities as verified by the State. PRWORA (Section 313).

- Grand Parent Liability: States must create procedures to make parents of a minor non-custodial parent with a child receiving welfare liable for paying child support. PRWORA (Section 373).

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2 The Effects of California Minimum Nurse Staffing Laws on Nurse Labor and Patient Mortality in Skilled Nursing Facilities

This paper investigates how a change in minimum nurse staffing regulation for California skilled nursing facilities (SNFs) affects nurse employment and how induced changes in nurse staffing affect patient mortality. In 2000, legislation increased the minimum nurse staffing standard and altered the calculation of nurse staffing, which created incentives to shift employment to lower skilled nurse labor. SNFs constrained by the new regulation increase absolute and relative hours worked by the lowest skilled type of nurse. Using this change in regulation to instrument for measured nurse staffing levels, I determine that the new staffing standard causes the number of patients discharged due to death per facility-year to decrease by 1.78, a 4.6% reduction from levels prior to 2000.

2.1 Introduction

This paper investigates how a change in minimum nurse staffing regulation in California skilled nursing facilities (SNFs) affects nurse employment and how regulation induced changes in nurse staffing affect patient mortality. A better understanding of regulations and quality of care in SNFs, otherwise known as nursing homes, is needed as the United States elderly population continues to grow. In this paper, I evaluate the effects from a change in minimum nurse staffing regulation and establish a causal relationship between nurse staffing and patient mortality.

Previous literature demonstrates that the impact of SNF staffing regulations on nurse staffing levels depends on the mandated threshold. Furthermore, these studies typically show that regulations and staffing levels are positively correlated with patient outcomes, but are unable to establish a causal relationship. Park and Stearns (2009) use national data to determine that state minimum nurse staffing laws cause modest increases in nurse employment in facilities with low initial staffing levels. They also find that state minimum nurse staffing regulations are associated with reductions in the use of restraints and the number of deficiencies. Mueller *et al.* (2006) categorize states as no standard, low standard, and high standard to establish a correlation between state minimum nurse staffing standards and nurse staffing levels. They find that states with high staffing standards have greater nurse staffing levels than states with low standards, and that staffing levels in states with low standards are not statistically different from staffing levels in states with no standards. Zhang and Grabowski (2004) evaluate the effects of the Nursing Home Reform Act (NHRA) of 1987, federal legislation that requires nursing homes certified under Medicare and Medicaid to maintain minimum levels of nurse staffing. Their results show that the passage of the NHRA is correlated with better care as measured by reductions in patients with pressure ulcers, physical restraints, and urinary catheters. Moreover, Zhang and Grabowski (2004) use a reduced form estimation to establish a positive correlation between registered nurse (RN) staffing, the highest skilled type of nurse, and outcomes of facilities with the lowest pre-NHRA staffing levels.

My study adds to the existing literature by estimating a causal relationship between nurse staffing and patient outcomes. There are many studies investigating the association between nurse staffing and patient outcomes in SNFs. These studies generally find a positive correlation between aggregate nurse staffing levels and patient outcomes (Spector and Takada, 1991; Harrington *et al.*, 2000; Schnelle *et al.*, 2004) and nurse skill mix and patient outcomes (Spector and Takada, 1991;

Cohen and Spector, 1996; Bleismer *et al.*, 1998). Related literature analyzing the relationship between nurse staffing and patient outcomes in acute care hospitals finds similar results (Aiken *et al.*, 2002; Needleman *et al.*, 2002; Unruh, 2003; Lang *et al.*, 2004; Kane *et al.*, 2007). While these studies provide evidence that higher nurse staffing and more skilled nurse staffing are associated with better care, they do not establish a causal relationship.

To my knowledge, there is only one study estimating a causal relationship between nurse staffing and outcomes in SNFs. A study by Konetzka *et al.* (2008) uses facility level variation in the implementation of the Medicare Prospective Payment Service system to estimate the causal effect of RN staffing levels on patient outcomes in SNFs. Using facility level data for five states, they determine that augmenting the absolute amount of RN hours decreases patient falls and urinary tract infections. In contrast, studies estimating causal relationships between nurse staffing and patient care in hospitals do not find staffing levels to significantly affect patient outcomes. Evans and Kim (2006) exploit short term patient admission shocks to determine that reductions in staffing per case do not significantly affect mortality in California hospitals, but do cause a decrease in patient length of stay and an increase in the likelihood of future readmission. Cook (2009) uses variation created by a regulation change to determine that nurse staffing levels do not significantly affect patient outcomes in California hospitals.

In this paper, I provide a direct analysis of the impact of minimum nurse staffing regulations while establishing a causal relationship between nurse staffing and patient outcomes in SNFs. I demonstrate that a change in California minimum nurse staffing regulation causes SNFs to increase nurse staffing levels. Then, I relate changes in nurse staffing induced by the new staffing regulation to patient outcomes using an instrumental variables approach.

The remainder of this paper is organized as follows. Section 2 provides background on SNFs and California minimum nurse staffing regulations, Section 3 describes the data, Section 4 outlines the economic model and predictions, Section 5 explains the empirical models and results, and Section 6 concludes.

2.2 Background

SNFs treat patients who need 24-hour medical care and employ three main types of nurses. Registered Nurses (RNs) are the highest educated and most expensive type of nurse. Individuals attend a 2-4 year nursing program at a post secondary institution to become a RN and 1 year of schooling to become a Licensed Vocational Nurse (LVN). LVNs are supervised by RNs and physicians and provide basic services such as taking vital signs and performing laboratory tests. RNs conduct diagnostic tasks which include recording patient symptoms and medical history, performing and interpreting results from diagnostic tests, and operating medical machinery. RNs are also in charge of dispensing certain medications that LVNs are not authorized to administer. RNs and LVNs are referred to as licensed nurses.

The third and least skilled type of nurse is a Certified Nurse Aide (CNA). To become a CNA in California, an individual must be at least 16 years of age, have at least 150 hours of training, and pass a competency exam. General duties of a CNA are to "answer patients' call lights; serve meals; make beds; and help patients eat, dress and bathe." (U.S. Bureau of Labor Statistics, 2006). Although nurse aides are not generally required to be certified, California staffing regulations only count CNAs when measuring staffing levels. CNAs work under the supervision of licensed nurses and physicians and are considered unlicensed nurses.

The California government measures staffing levels using hours per resident day (HPRD). HPRD is calculated by dividing total nurse hours[1] by total patient days.[2] Minimum HPRD is not adjusted for patient acuity and applies to the entire patient population regardless of source of payment. Prior to 2000, total nurse hours were defined as CNA hours plus two times the sum of RN and LVN hours. On January 1, 2000, the minimum HPRD was increased from 3.0 to 3.2 and the

[1]Total nurse hours include permanent and temporary nurse employees. If a SNF has less than 60 beds, then nurse management hours are added to total nurse hours. Approximately 25% of SNFs in the sample have less than 60 beds. Total nurse hours also include geriatric nurse hours. While nurse management and geriatric nurse hours are included in the HPRD calculation, their values are negligible and are excluded from the empirical analysis.

[2]Total patient days is defined as the number of days patients spend in a SNF which includes the day admitted, but not the day discharged.

double counting of RN and LVN hours was eliminated. Therefore, this change in minimum nurse staffing regulation increased the relative weight of CNA labor while mandating an aggregate increase in total nurse staffing. With a trade off between more total hours and fewer skilled hours, it is unclear how patient outcomes will be affected. I use this change in regulation to provide exogenous variation in staffing levels to identify a causal relationship between HPRD and patient mortality.

The new regulation was formally implemented on April 1, 2000. To determine compliance, annual inspectors randomly choose two weeks out of the year to measure HPRD. Hence, SNFs do not have to meet the 3.2 threshold in real time, but only need to average 3.2 HPRD within every two week period. Violations of the regulation result in fines and possible license suspension. Despite this lack of supervision, the percentage of facilities averaging staffing levels above the 2000 threshold steadily increased from 25% in 1999 to 65% in 2002, demonstrating that the new regulation influences the amount of nurses employed by SNFs. To my knowledge, there are no other changes in SNF regulations during this time period.[3] Federal law requires SNFs to employ 1 RN for 8 consecutive hours, 1 RN or LVN for the remaining two 8 hour shifts, and 0.30 licensed HPRD for facilities with more than 100 residents.[4] In practice, average licensed HPRD in California SNFs is higher than the federal regulation, perhaps because patient caseload necessitates more skilled labor and additional licensed staff is needed to supervise CNAs.

The Balanced Budget Act (BBA) of 1997 reformed Medicare coverage the same year that the California regulation went into effect. Under the BBA, Medicare reimbursement for SNF care changed from fee-for-service to a Prospective Payment Service (PPS) system in 2000.[5] Grabowski and Gruber (2007) show that

[3]In 2004, California legislation changed minimum nurse staffing laws in acute care hospitals. I restrict the data to years 1995-2002 to rule out possible interactions between hospital and SNF nurse labor markets. Preliminary studies determine that the change in minimum nurse staffing regulation for acute care hospitals cause average nurse to patient ratios to decrease; however, the effect on patient outcomes is mixed (Donaldson *et al.*, 2005; Bolton *et al.*, 2007; Cook, 2009).

[4]SNFs must also have one RN who is a full time director of nursing. If a SNF has fewer than 60 residents, then this director of nurse may also be a charge nurse.

[5]The BBA also imposed the PPS reimbursement system for Medicare coverage of home health care. Even though SNF and home health care are partial substitutes for each other, the degree to which the two forms of care are substitutes depends on physician diagnosis. To qualify for Medicare coverage of SNF care, a patient must be hospitalized for at least 3 days prior and must be recommended for SNF entry by a physician. To qualify for Medicare coverage of home health

nursing home demand is inelastic to public program generosity, namely Medicaid, suggesting that the BBA should not impact SNF entry from Medicare patients. The PPS system reimburses SNFs with fixed payments based on patient case mix, providing SNFs with incentives to characterize admitted patients into categories that provide higher reimbursement and to provide less expensive care (White, 2003). Selectively admitting patients that require less costly treatments will potentially cause patient acuity to decline. If the BBA reduces patient acuity, then my results may over estimate any positive impact from increasing staffing levels. Although the BBA calls into question the validity of using the California regulation as an instrument, only a small percentage of California SNF patients pay with Medicare, less than 10%. Therefore, the population at-risk of being affected by both the BBA and California regulation is small.

2.3 Data

I use annual SNF data from the California Office of Statewide Health Planning and Development (OSHPD) for years 1995 through 2002. To achieve comparable statistics across years, I restrict the data to the 812 state certified SNFs surveyed each year.[6] The new regulation is predicted to have an impact on nurse labor because 75% of SNFs will average less than 3.2 HPRD if they do not adjust nurse staffing levels in 2000.

Even though surveyors measure HPRD using two weeks of data, I calculate HPRD using annual data and refer to a facility as being in compliance if the annual average is above the minimum threshold. HPRD calculated under the 2000 regulation will be called HPRD2000. I define High Staffing SNFs as SNFs with 1999 HPRD2000 values *above or equal to* the 3.2 threshold, and Low Staffing SNFs as SNFs with 1999 HPRD2000 values *below* the 3.2 threshold. High Staffing SNFs

care, a patient must require intermittent skilled care or must be confined to stay inside a home as prescribed by a physician. Given that eligibility for Medicare coverage does not change, the BBA is not expected to affect the distribution of patients using SNF and home health care.

[6]Facilities that employ zero nurse hours, report zero patients, or experience inconsistencies in data are also dropped. Additionally, facilities that do not pass the pre-2000 regulation in 1999 are dropped. These facilities do not abide by the existing regulation and are not expected to achieve the 2000 standard. These restrictions cause the sample size to drop from 912 to 812.

are facilities that already meet the 2000 standard prior to its implementation, whereas Low Staffing SNFs are facilities that will average less than 3.2 HPRD if they do not change staffing levels.

Table 2.1 contains summary statistics by pre-law (1995-1999) and post-law (2000-2002) time periods and reveals distinct differences between Low and High Staffing SNFs.[7] Pre-law HPRD2000 is about 30% higher for High Staffing SNFs than Low Staffing SNFs. High Staffing SNFs also have approximately 60% more RNs and 25% more LVNs and CNAs than Low Staffing SNFs. High Staffing SNFs are smaller in size and have a lower fraction of Black and Hispanic patients. Compared to Low Staffing SNFs, High Staffing SNFs have fewer patients paying with Medi-Cal, California's Medicaid program, and more patients paying for their own services, demonstrating that these facilities cater to higher income patients.

2.4 Economic Model

To provide predictions on how the 2000 minimum nurse staffing regulation affects the employment of each type of nurse, I use a model of firm profit maximization similar to the ones employed by Nyman (1985) and Gertler (1989, 1992). Unlike hospitals, most nursing homes are for-profit institutions. In the California data, 85% of all SNFs and 95% of Low Staffing SNFs are for-profit.[8] Because Cal-

[7]The coefficient of variation is higher for High Staffing SNFs than Low Staffing SNFs likely because the sample size of High Staffing SNFs is small, causing average statistics to be less precise. Comparing pre-law and post-law averages reveals an increase in aggregate nurse staffing levels for High Staffing SNFs. This increase in average nurse staffing levels is partially explained by an increase in LVN HPRD which reflects a secular shift in licensed nurse hours to be discussed in Section 5.1.2. There is also an average increase in CNA HPRD. In Section 5, I do not find CNA staffing in High Staffing SNFs to be significantly affected by the new regulation, suggesting that this change is explained by other covariates.

[8]For-profit SNFs and SNFs affiliated with a chain operate differently than non-profit facilities (Davis, 1993; Spector *et al.*, 1998; O'Neill *et al.*, 2003). In particular, proprietary SNFs typically have lower average costs and lower quality of care than non-profit facilities. In this study, the main sample of interest is Low Staffing SNFs, which are 95% for-profit. SNFs that are part of a chain may be more efficient in terms of reallocating resources to meet the new regulation. Furthermore, if a system chain has standardized training and management, then the regulation change may have a more beneficial impact on patient outcomes for SNFs affiliated with a chain (Kamimura *et al.*, 2007). Unfortunately, I do not observe system affiliation in the California OSHPD data. However, including facility fixed effects should account for both chain affiliation and for-profit status since these are generally time invariant characteristics.

ifornia minimum nurse staffing requirements weight licensed nurse hours equally, I consider a simplified case with two types of nurses, licensed nurses L and unlicensed nurses U. Each firm produces health services S and the production function F depends on quality Q and nurse staffing levels L and U. Price of health services p is increasing in quality Q.

With heterogeneity in patients' ability to pay for health services, SNFs operate at different levels of quality to capture each market share. If SNF quality depends on nurse staffing levels and nurse skill mix, then the data provide evidence that SNFs produce at different quality levels because nurse staffing varies across facilities. For simplicity, I assume that each SNF produces health services within a particular quality segment. Each SNF chooses the combination of L and U that maximizes profits, and facilities enter and exit until profits are equalized across each quality segment. In this framework, nurse labor is the only cost to a SNF. I assume that wages remain constant[9] and wages, w_i, and marginal product, F_i, adhere to the following inequality constraints: 1) $w_L > w_U$ and 2) $F_L > F_U$.

Before the minimum staffing constraint changes, SNFs within each quality segment maximize profits subject to the existing minimum staffing requirement:

$$max_{L,U} \quad p(\overline{Q})F(L,U,\overline{Q}) - C(L,U)$$

subject to

$$a_1 L + b_1 U \geq K_1$$

where a_1 and b_1 are minimum nurse staffing weights and K_1 is a minimum threshold. The cost function is a linear function, $C(L,U) = w_L L + w_U U$. I assume that the relative ratio, $\frac{a_1}{b_1}$, is less than the relative ratio, $\frac{w_L}{w_U}$, which is supported empirically in California. This causes the slope of the minimum cost line to be steeper than the slope of the minimum staffing constraint. where a_1 and b_1 are minimum

[9]Based on the model, staffing levels for each type of nurse will depend on relative wage, which may differ across facilities. However, including wages as explanatory variables in the empirical estimation would cause the estimates to suffer from endogeneity because wages themselves may be affected by relative demand. For simplicity, I assume that relative wages remain constant. According to the U.S. Bureau of Labor Statistics, California nurse wages remain constant with hourly RN wage roughly equaling 1.5 times the hourly wage of LVNs and 2.8 times the hourly wage of CNAs for years before and after the regulation change in 2000.

nurse staffing weights and K_1 is a minimum threshold. The cost function is a linear function, $C(L,U) = w_L L + w_U U$. I assume that the relative ratio, $\frac{a_1}{b_1}$, is less than the relative ratio, $\frac{w_L}{w_U}$, which is supported empirically in California. This causes the slope of the minimum cost line to be steeper than the slope of the minimum staffing constraint.

When the new regulation is imposed, SNFs with staffing levels below the new constraint will fall out of compliance unless they alter staffing levels as depicted in Figure 2.1. In the data, I call these facilities Low Staffing SNFs. SNFs with staffing levels above the new constraint do not need to change staffing levels to remain in compliance. These facilities are called High Staffing SNFs. Given that High Staffing SNFs already meet the new staffing constraint, aggregate nurse staffing levels are predicted to be unaffected for this group.

Although the production isoquants and cost minimizing expenditure lines may vary across quality segments, I simplify the graphical analysis in Figure 2.1 by depicting the isoquant associated with the profit maximizing level of production for Low Staffing SNFs in one particular quality segment. Prior to the policy change, these SNFs optimize profits at point (L1,U1). When the new regulation is imposed, SNFs will re-maximize profits subject to the new constraint:

$$a_2 L + b_2 U \geq K_2$$

I examine the case where the new staffing regulation increases the minimum threshold ($K_2 > K_1$) and decreases the relative weight of the more expensive type of nurse ($\frac{a_2}{b_2} < \frac{a_1}{b_1}$), which is what occurs in California with the 2000 regulation change. Again, I assume that the relative ratio, $\frac{a_2}{b_2}$, is less than the relative ratio, $\frac{w_L}{w_U}$.

To understand the implications of the staffing law change, refer to Figure 2.2. Holding production and quality fixed, SNFs will choose the combination of L and U that passes the new staffing constraint at minimum cost. This outcome (L2,U2) occurs at the intersection of the production isoquant and the new staffing constraint. The graphical analysis demonstrates that the 2000 minimum nurse staffing law will cause an increase in unlicensed nurse hours, a decrease in licensed nurse hours, and an increase in measured HPRD. If SNF production is allowed to vary, then SNFs will scale back production to operate along the original minimum cost line at point (L3,U3). At (L3,U3), Low Staffing SNFs experience even larger increases in unlicensed nurse hours and larger decreases in licensed nurse hours than is the case with fixed output.

The trade off between licensed and unlicensed nurse hours is more pronounced for Low Staffing SNFs that are initially farther from reaching the new minimum K_2. I empirically test for differential effects based on how much a Low Staffing SNF needs to increase its nurse staffing levels to meet the 3.2 threshold. Using 1999 data, I define a time invariant variable called Threshold Gap which equals 3.2 - HPRD2000. Facilities with higher Threshold Gap values are predicted to fail the 2000 standard by a greater margin if nurse staffing is left unchanged. This can be seen graphically in Figure 2.2 as the perpendicular distance between (L1,U1) and the new staffing constraint. Graphically, Figure 2.2 shows that a more positive Threshold Gap yields a larger perpendicular distance between (L1,U1) and the new staffing constraint, causing there to be more substitution of unlicensed nurse hours for licensed nurse hours.

The pair-wise conclusions discussed above hold when extending the model to the three type case because relative nurse wages are assumed to remain constant. The 2000 regulation will cause the most expensive type of nurse, RNs, to decrease, the cheapest type of nurse, CNAs, to increase, and have an ambiguous effect on LVN staffing. The economic model predicts that Low Staffing SNFs will increase total nurse employment and substitute labor away from RNs to CNAs. Moreover, this model predicts that High Staffing SNFs will not be affected by the new minimum staffing constraint.

In practice, Low Staffing SNFs may choose to move into different quality segments in response to the new regulation. When quality is allowed to vary, the predictions on nurse staffing are the same as before. The effect of the new standard on quality in Low Staffing SNFs is ambiguous because we do not know the full production function. If quality only depends on nurse staffing levels, then quality will increase or decrease depending on which factor affects it more, total nurse labor or nurse skill mix. If other factors besides nurse staffing affect quality, then Low Staffing SNFs may adjust quality to counteract any reduction in profits caused by changes in nurse staffing. To test how the 2000 minimum nurse staffing standard affects quality in Low Staffing SNFs, I estimate how changes in aggregate nurse staffing levels induced by the new regulation affect the number of patients discharged due to death per facility-year.

2.5 Empirical Analysis

I use a SNF fixed trends estimation[10] to test the predictions of the economic model separately for Low and High Staffing SNFs. It might seem natural to implement a differences-in-differences estimation using the sample of High Staffing SNFs as a control group. However, Low and High Staffing SNFs differ across observable characteristics and exhibit different pre-law nurse staffing trends, suggesting that a separate analysis is more appropriate.

2.5.1 Effect on Nurse Staffing

To examine how the 2000 regulation affects nurse staffing, I estimate the following model:[11]

$$Y_{it} \quad = \quad A_i + t + A_i t + \omega Post_t + \Gamma X_{it} + \pi C_{it} + \eta_{it}$$

where Y_{it} is equal to the nurse staffing outcome of interest for facility i in year t. X_{it} contains SNF patient demographics including race, age, gender, and method of payment. SNF fixed effects account for omitted time invariant facility characteristics such as a propensity to treat patients with severe illnesses and system affiliation. SNF fixed trends account for facility specific linear trends in patient

[10]Unreported results show that the empirical results are insensitive to the addition of county fixed trends and county quadratic trends.

[11]To understand how using different specifications affect the coefficient estimates, I started with a basic linear regression and systematically added SNF fixed effects, a linear trend, and linear SNF fixed trends. Compared to the baseline model, the inclusion of SNF fixed effects causes the coefficient on $Post_t$ to reduce by roughly 25% when estimating its impact on nurse staffing levels. Adding a linear trend and linear SNF fixed trends does not significantly change the coefficient estimates obtained from the SNF fixed effects model.

case mix or other variables not included in X_{it}. C_{it} is vector of county level characteristics, including the poverty rate and unemployment rate. $Post_t$ equals one when $t \geq 2000$ and zero otherwise. The coefficient estimate of interest, ω, measures the effect of the new minimum nurse staffing standard on Y_{it}.

For Low Staffing SNFs, I also test for differential impacts by including $Post_t * ThresholdGap_i$ in a separate specification, where Threshold Gap is normalized to have mean zero and standard deviation of 1. The coefficient estimate on the interaction term is the differential impact of the law change on a Low Staffing SNF with a Threshold Gap value that is one standard deviation above the mean.

Low Staffing SNFs

Results in Table 2.2 reveal that Low Staffing SNFs respond to the change in minimum nurse staffing standard by increasing overall nurse hours.[12,13] The new regulation causes HPRD2000 to increase by an average of 0.29, a 10% increase from pre-2000 levels. Statistically significant at the 1% level, coefficient estimates show that the new standard causes RN HPRD to decrease by 0.03, LVN HPRD to increase by 0.066, and CNA HPRD to increase by 0.26. These results demonstrate that the increase in total nurse staffing mainly comes from CNA employment. The new staffing standard causes licensed HPRD to increase by 0.035 and the percentage of licensed hours to decrease by 1.5 percentage points. Therefore, even though the regulation causes total licensed nurse hours to increase, it also causes a decrease in nurse skill mix.

If CNAs were perfect substitutes for licensed nurses, then there would be a decrease in total licensed hours. A small increase in licensed nurse hours reveals that licensed nurses perform different tasks than CNAs as described in Section 2.

[12]HPRD2000 is calculated by dividing total nurse hours by patient days. To ensure that increases in HPRD2000 are caused by changes in nurse hours, I estimate the effect of the new regulation on total nurse hours and patient days separately using equation (5.1). Unreported results show that increases in HPRD2000 are coming from statistically significant changes in nurse hours and not from patient days.

[13]To determine whether or not Low Staffing SNFs increase staffing levels prior to or after the regulation change in 2000, I re-estimated the model adding in the lag and lead of the $Post_t$ variable. Unreported results demonstrate that Low Staffing SNFs begin increasing staffing levels in 2000, with more modest increases occurring in 2001 and 2002.

Furthermore, SNFs are unable to decrease licensed nurse hours to compensate for the costs of hiring more CNAs, suggesting that SNFs operate at some optimal level of licensed nurse hours. Instead of being curved as in Figures 2.1 and 2.2, the production isoquants asymptote to some minimum level of licensed nurse hours. This optimal level could be dictated by patient case mix and supervision constraints.

Results in Table 2.3 show that a one standard deviation increase in Threshold Gap yields an additional 0.043 increase in CNA HPRD, meaning that Low Staffing SNFs which need to increase their staffing levels the most to achieve the new standard do so by increasing CNA labor. Although changes in RN and LVN hours are predicted to vary by Threshold Gap, I do not find statistically significant differential effects on RN, LVN, or total licensed nurse hours. These results provide additional evidence that facilities maintain a minimum level of licensed nurse staffing. Moreover, these changes in RN and LVN hours could be caused by a secular shift in licensed staffing, which I discuss in the following section.

High Staffing SNFs

The new staffing standard does not have a statistically significant impact on aggregate nurse staffing levels among High Staffing SNFs. However, results reveal that the new regulation causes RN HPRD to decrease by 0.056 and LVN HPRD to increase by 0.046. The new staffing standard does not significantly affect licensed HPRD or the percentage of licensed nurse hours, implying that changes in RN and LVN hours offset each other.

Finding any significant effect on nurse staffing is perplexing. High Staffing SNFs already comply with the 2000 standard prior to its passage and should be unaffected by the policy change. One possible explanation is that there is a secular shift in RN and LVN labor occurring at the same time as the policy change. In particular, there is a RN shortage during this time period (Buerhaus *et al.*, 2003) that may explain the decline in RN hours in both Low and High Staffing SNFs. Since the increase in LVN hours compensates for the decline in RN hours, SNFs appear to operate at some optimal amount of licensed hours.

Results for High Staffing SNFs suggest that the changes in RN and LVN staffing in Low Staffing SNFs could be a result of a secular shift in licensed staffing and should be interpreted with caution. The regulation change does not significantly affect CNA labor in High Staffing SNFs, providing support that increases in CNA labor in Low Staffing SNFs are caused by the new staffing standard.[14,15] Because there is no significant change in total nurse staffing levels in High Staffing SNFs, I restrict the second stage analysis of how staffing levels affect patient mortality to the sample of Low Staffing facilities.

2.5.2 Patient Mortality

In this section, I instrument for measured nurse staffing levels using $Post_t$ to estimate the effect of nurse staffing on patient mortality for Low Staffing SNFs.[16] Ordinary Least Squares (OLS) estimation is believed to be biased because facilities can choose staffing levels and nurse skill mix based on patient acuity. Konetzka *et al.* (2008) speculate that facilities with sicker patients likely choose higher staffing levels, implying that the true effect of staffing on outcomes is under estimated by OLS. By using the 2000 regulation change to instrument for nurse staffing levels, I assume that the only way the new regulation affects patient mortality is through changes in HPRD2000. Put differently, I assume that changes in nurse staffing are not caused by changes in patient acuity, which may or may not be true. I acknowledge that a weakness of the Instrumental Variables (IV) strategy is a possible correlation between policy induced changes in nurse staffing and patient

[14]To provide further evidence that staffing increases in Low Staffing SNFs are due to the regulation change, I check to see whether neighboring states experience similar changes. Studies conducted by the American Health Care Association (1997, 2004) show that average full time equivalent nursing staff per bed and average full time equivalent nursing staff per resident decrease in Arizona and Nevada and increase in California between 1997 and 2003. Although these measures do not explicitly measure HPRD or include differentiation between types of nurses, they provide some evidence that the increase in nurse staffing in Low Staffing SNFs is unique to California SNFs.

[15]I examine how nurse staffing levels change in California hospitals during the time period when the California SNF regulation was implemented. Using California hospital data from OSHPD, I determine that increases in nurse HPRD only occur in California SNFs and not in California hospitals. This exercise provides additional evidence that the California regulation only affects nurse staffing in SNFs and does not reflect general changes in the nurse labor market.

[16]Results when instrumenting for HPRD using the pre-2000 calculation are analogous to the results reported in this section. Therefore, results are not driven by using HPRD2000 as the measure of nurse staffing levels.

acuity. Consequently, I control for patient acuity as much as possible by including SNF fixed effects, SNF fixed trends, and SNF level patient demographics.

The 2SLS estimation is:[17,18]

$$D_{it} = A_i + t + A_i t + \omega \widehat{HPRD2000}_{it} + \Gamma X_{it} + \pi C_{it} + \epsilon_{it}$$

where D_{it} is the number of patients discharged due to death in SNF i in year t and $\widehat{HPRD2000}_{it}$ are fitted values of HPRD2000 obtained by estimating equation (5.1). Because I have one instrumental variable, I am unable to estimate how each type of nurse labor affects mortality. Instead, I estimate how aggregate changes in nurse staffing, as measured by HPRD2000, impact mortality while keeping in mind that these changes are largely driven by increases in CNA nurse hours.

Results in Table 2.5 show that a one unit increase in HPRD2000 causes a 6.14 decline in the number of patients discharged due to death per facility-year.[19,20] Thus, regulation induced changes in HPRD2000 cause the number of patients discharged due to death to decrease by 1.78 per facility-year,[21] a 4.6% reduction from pre-law levels. To demonstrate that the reduction in deaths is not a result of an overall shift in patient mortality trends, I conduct a reduced form estimation of the effect of $Post_t$ on patient mortality in High Staffing SNFs. The reduced form estimation yields insignificant results, demonstrating that the regulation change

[17] The first stage partial F-statistic on $Post_t$ is 170 and passes Staiger and Stock's (1997) rule of thumb.

[18] Similar to the first stage analysis, I systematically add SNF fixed effects, a linear trend, and linear SNF fixed trends to the baseline model to understand how different specifications affect the Instrumental Variable (IV) estimates. Adding SNF fixed effects to the baseline IV model reduces the impact of HPRD2000 on the number of patients discharged due to death per facility-year by 20% while the addition of a linear trend and SNF fixed trends does not significantly alter the SNF fixed effects results.

[19] Estimating the causal impact of nurse staffing on the number of patients discharged due to death per facility-year using the $Post_t$ variable and its interaction with Threshold Gap as instruments yields analogous, but less precise results.

[20] For the entire sample of SNFs, I use $Post_t$ and its interaction with Threshold Gap to instrument for aggregate nurse staffing levels. Similar to the results presented in the main empirical section, higher nurse staffing levels cause a decrease in the number of patients discharged due to death per facility-year. Since I lose precision when implementing the IV estimation on the whole sample with these two instruments, I restrict the main analysis to Low Staffing SNFs and instrument staffing levels with $Post_t$ only.

[21] 1.78 number of patients discharged due to death per facility-year is obtained by multiplying 6.14 with 0.29, the coefficient estimate on $Post_t$ from the first stage.

is not associated with mortality in High Staffing SNFs and that IV results for Low Staffing SNFs are not caused by a general shift in patient mortality trends. Due to concurrent changes in Medicare reimbursement, the $Post_t$ variable may partially capture changes from the BBA as discussed in Section 2. Hence, the 1.78 decrease in patient deaths per facility-year should be viewed as an upper bound of the impact from regulation induced changes in nurse staffing.

Results in Table 2.5 show that a one unit increase in HPRD2000 causes a 6.14 decline in the number of patients discharged due to death per facility-year.[19,20] Thus, regulation induced changes in HPRD2000 cause the number of patients discharged due to death to decrease by 1.78 per facility-year,[21] a 4.6% reduction from pre-law levels. To demonstrate that the reduction in deaths is not a result of an overall shift in patient mortality trends, I conduct a reduced form estimation of the effect of $Post_t$ on patient mortality in High Staffing SNFs. The reduced form estimation yields insignificant results, demonstrating that the regulation change is not associated with mortality in High Staffing SNFs and that IV results for Low Staffing SNFs are not caused by a general shift in patient mortality trends. Due to concurrent changes in Medicare reimbursement, the $Post_t$ variable may partially capture changes from the BBA as discussed in Section 2. Hence, the 1.78 decrease in patient deaths per facility-year should be viewed as an upper bound of the impact from regulation induced changes in nurse staffing.

This study is the first to my knowledge to establish a significant causal relationship between aggregate nurse staffing levels and patient mortality in SNFs. Studies by Evans and Kim (2006) and Cook (2009) do not find nurse staffing levels to have a significant effect on patient mortality in California hospitals, but

[19] Estimating the causal impact of nurse staffing on the number of patients discharged due to death per facility-year using the $Post_t$ variable and its interaction with Threshold Gap as instruments yields analogous, but less precise results.

[20] For the entire sample of SNFs, I use $Post_t$ and its interaction with Threshold Gap to instrument for aggregate nurse staffing levels. Similar to the results presented in the main empirical section, higher nurse staffing levels cause a decrease in the number of patients discharged due to death per facility-year. Since I lose precision when implementing the IV estimation on the whole sample with these two instruments, I restrict the main analysis to Low Staffing SNFs and instrument staffing levels with $Post_t$ only.

[21] 1.78 number of patients discharged due to death per facility-year is obtained by multiplying 6.14 with 0.29, the coefficient estimate on $Post_t$ from the first stage.

this discrepancy could be because SNFs and hospitals have different production functions. The reduction in mortality supports a study by Park and Stearns (2009) that reports a positive correlation between staffing and quality of care. In addition, this result does not necessarily support or disprove studies showing that RN staffing improves quality of care (Zhang and Grabowski, 2004; Konetzka *et al.*, 2008) because the estimation strategy only identifies the impact of total nurse hours. Since this study uses the California regulation change to instrument for nurse staffing levels, the IV results represent a local average treatment effect in which the negative, significant effect of nurse staffing on patient mortality applies to Low Staffing SNFs that experience regulation induced increases in HPRD2000.

A reduction in SNF mortality suggests that the benefits from increasing overall staffing levels outweighs any adverse effect from reducing nurse skill mix. Higher CNA hours are associated with providing better daily care processes (Schnelle *et al.*, 2001) and may subsequently decrease patient mortality. Moreover, if facilities are already operating at an adequate level of licensed nurse staffing, then increasing unlicensed nurses might improve patient outcomes (Lang *et al.*, 2004). For example, CNAs may ease the workload of licensed nurses and increase the likelihood of detecting a patient in distress. In addition, an overall increase in staffing levels will reduce nurse-patient caseload and might improve outcomes by decreasing nurse burnout and dissatisfaction (Aiken *et al.*, 2002).

Although I determine that increasing staffing levels reduces patient deaths within a facility, interpreting this result as a positive outcome requires caution because this mortality measure is not adjusted for life expectancy. Therefore, the value of reducing mortality depends on how long patients live after being discharged and the quality of their health. On average, 25% of patients are discharged due to death and the rest are discharged home or to another medical facility. Unreported results find staffing levels to have a positive, but insignificant impact on the fraction discharged to other facilities or to their home. If the reduction in mortality comes from an increase in patients being sent to hospitals or from an increase in patients discharged home who pass away shortly after, then this result should not be viewed as an improvement in patient care. Unfortunately, the data do not follow patients after discharge so I am unable to make assertions on a patient's quality of life after he/she leaves a SNF.

2.6 Discussion and Conclusion

In this paper, I demonstrate that Low Staffing SNFs respond to the 2000 minimum nurse staffing regulation by increasing total staffing by 10%, which is mostly caused by augmenting CNA nurse labor. Changes in RN and LVN labor in both Low and High Staffing SNFs suggest that there could be a secular shift in licensed nurse hours. Low Staffing SNFs are unable to decrease licensed nurse hours to compensate for costs from increasing CNA hours, providing evidence that SNFs maintain some minimum level of licensed nurse hours. Using the new staffing standard to instrument for measured nurse staffing levels, I establish a positive causal relationship between nurse staffing levels and patient mortality. Changes in nurse staffing induced by the new minimum staffing regulation decrease the number of patients discharged due to death by 1.78 per facility-year, a 4.6% reduction in patient mortality.

These results are short-term because this study only includes 3 years of post law data. As of 2002, 57% of Low Staffing SNFs meet the 2000 standard,[22] implying that the final impact from the new regulation may not be fully realized. However, these results suggest that minimum nurse staffing regulations are effective in decreasing nurse caseload. More CNA hours might yield better outcomes by improving daily care, reducing the workload of licensed nurses, and raising the likelihood of detecting patients in distress.

Patient mortality is arguably the most extreme measure of quality of care. As discussed in Section 5.2, interpreting estimates of patient mortality requires caution because the health status of patients upon discharge is not observed. The value of each life saved depends on how long a patient lives and his/her quality of life after leaving a facility. Ideally, I would also estimate the causal relationship between nurse staffing and less extreme measures such as the incidence of falls or

[22] Harrington and O'Meara (2006) estimate that 73% of California SNFs are in compliance with the 2000 standard in 2003.

pressure sores because these outcomes would allow me to tease out more subtle quality differences. Future studies might investigate the impact of minimum nurse staffing regulation on less extreme outcome measures and consider the effects of minimum staffing standards on non-labor inputs.

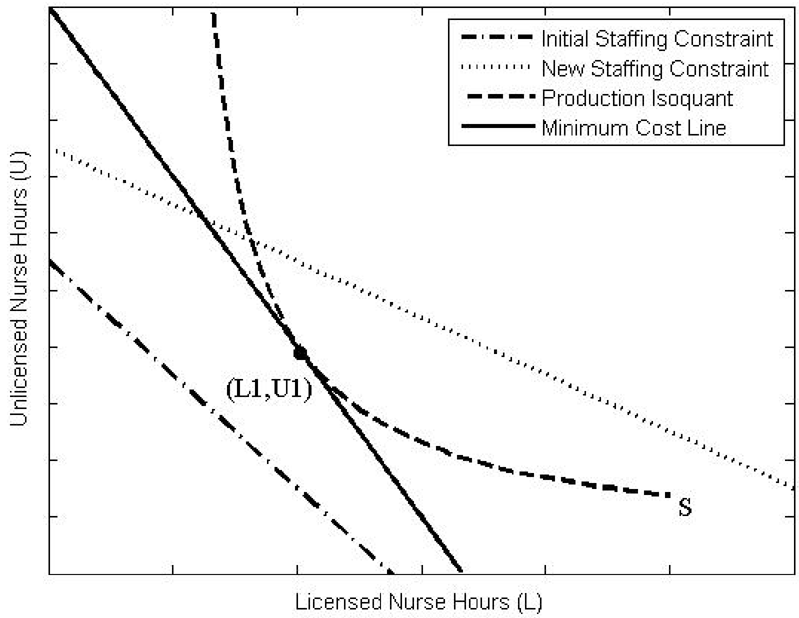

Figure 2.1: Profit maximization and staffing constraints

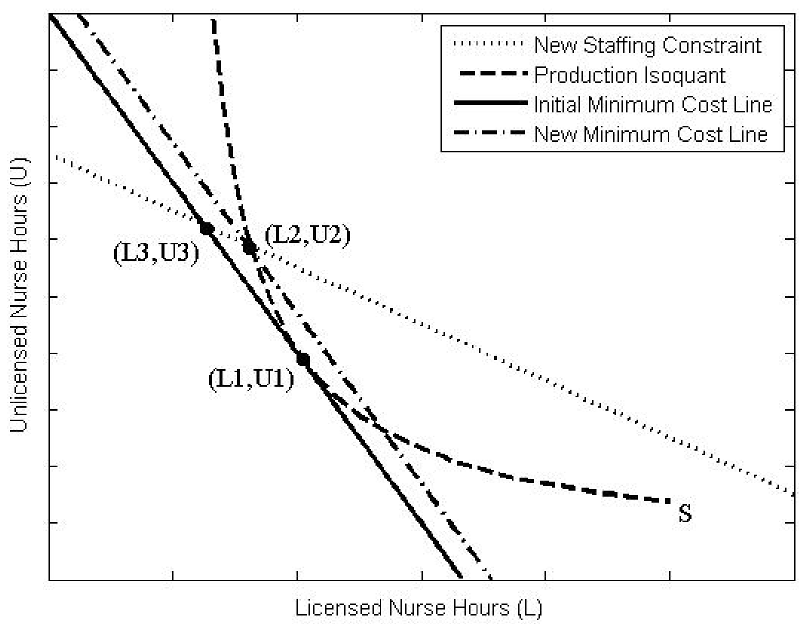

Figure 2.2: Equilibrium for Low Staffing SNFs under the new staffing constraint

Table 2.1: Summary Statistics

Variable	Low Staffing SNFs Pre-law 1995-1999		Low Staffing SNFs Post-law 2000-2002		High Staffing SNFs Pre-law 1995-1999		High Staffing SNFs Post-law 2000-2002	
HPRD2000	2.819	(0.602)	3.122	(0.353)	3.693	(1.011)	3.842	(1.14)
RN HPRD	0.297	(0.216)	0.281	(0.152)	0.487	(0.491)	0.456	(0.45)
LVN HPRD	0.507	(0.209)	0.581	(0.178)	0.633	(0.316)	0.709	(0.35)
CNA HPRD	1.987	(0.383)	2.236	(0.303)	2.502	(0.528)	2.608	(0.701)
% Licensed Hours	0.285	(0.058)	0.276	(0.054)	0.297	(0.069)	0.299	(0.073)
% Male Patients	0.299	(0.124)	0.318	(0.13)	0.249	(0.131)	0.259	(0.121)
% Patients Age 85+	0.409	(0.167)	0.38	(0.172)	0.538	(0.207)	0.526	(0.222)
% Black Patients	0.109	(0.173)	0.119	(0.171)	0.05	(0.117)	0.054	(0.117)
% Hispanic Patients	0.101	(0.11)	0.122	(0.125)	0.049	(0.08)	0.062	(0.095)
% Paid by Medi-Cal	0.689	(0.221)	0.694	(0.222)	0.395	(0.303)	0.413	(0.312)
% Paid by Medicare	0.074	(0.083)	0.08	(0.086)	0.077	(0.126)	0.082	(0.122)
% Paid by Self	0.177	(0.188)	0.146	(0.174)	0.42	(0.322)	0.381	(0.33)
# of Licensed Beds	103.679	(47.137)	103.618	(47.092)	88.133	(53.894)	88.163	(53.765)
Occupancy Rate	92.886	(26.431)	89.949	(22.063)	89.4	(29.441)	87.825	(17.92)
# Patient Deaths	38.325	(28.894)	36.387	(28.423)	41.901	(36.169)	40.402	(44.608)
County Unemployment Rate	7.213	(3.313)	6.034	(1.936)	6.402	(3.303)	5.679	(1.958)
County Poverty Rate	16.331	(5.024)	13.654	(4.044)	14.913	(5.373)	12.471	(4.315)
Observations	3060		1836		1000		600	

Notes: HPRD2000 = hours per resident day under the 2000 calculation, RN HPRD = registered nurse hours per resident day, LVN HPRD = licensed vocational nurse hours per resident day, and CNA HPRD = certified nurse aide hours per resident day. Low Staffing SNFs are facilities with HPRD2000 values less than 3.2 in 1999. High Staffing SNFs are facilities with HPRD2000 values greater than or equal to 3.2 in 1999. Standard errors reported in parentheses.

Table 2.2: Nurse Staffing Results for Low Staffing SNFs

Variables	HPRD2000	RN HPRD	LVN HPRD	CNA HPRD	Licensed HPRD	% Licensed
Post	0.290**	-0.031**	0.066**	0.260**	0.035**	-0.015**
	(0.022)	(0.007)	(0.009)	(0.017)	(0.010)	(0.002)
% Male Patients	0.153	0.027	0.047	0.067	0.075	0.002
	(0.092)	(0.035)	(0.039)	(0.076)	(0.052)	(0.013)
% Patients Age 85+	-0.069	0.027	-0.118**	0.024	-0.091**	-0.020*
	(0.065)	(0.025)	(0.031)	(0.052)	(0.034)	(0.009)
% Black Patients	0.093	-0.051	0.155**	-0.005	0.104*	0.021
	(0.101)	(0.049)	(0.051)	(0.084)	(0.053)	(0.015)
% Hispanic Patients	0.319*	-0.025	0.074	0.255*	0.049	0.002
	(0.151)	(0.045)	(0.055)	(0.116)	(0.061)	(0.014)
% Paid by Medi-Cal	0.065	-0.069	-0.038	0.204	-0.107	-0.039*
	(0.146)	(0.040)	(0.048)	(0.134)	(0.076)	(0.016)
% Paid by Medicare	0.007	0.014	-0.117*	0.138	-0.103	-0.026
	(0.169)	(0.051)	(0.056)	(0.140)	(0.086)	(0.017)
Constant	2.738**	0.439**	0.484**	1.727**	0.923**	0.336**
	(0.179)	(0.054)	(0.109)	(0.148)	(0.138)	(0.036)
Observations	4896	4896	4896	4896	4896	4896
Number of SNFs	612	612	612	612	612	612
R-squared	0.337	0.259	0.255	0.331	0.293	0.156

Notes: * Significant at the 5% level. ** Significant at the 1% level. SNF clustered robust standard errors reported in parentheses. SNF fixed effects and fixed trends included. Coefficient estimates on selected explanatory variables not reported for brevity. RN = registered nurse, LVN = licensed vocational nurse, CNA = certified nurse aide, and HPRD2000 = hours per resident day under the 2000 calculation. Post = 1 if t=2000-2002 and = 0 for t<2000. Low Staffing SNFs are facilities with HPRD2000 values less than 3.2 in 1999.

Table 2.3: Nurse Staffing Results for Low Staffing SNFs By Threshold Gap

VVariables	HPRD2000	RN HPRD	LVN HPRD	CNA HPRD	Licensed HPRD	% Licensed
Post	0.296**	-0.030**	0.062**	0.269**	0.032**	-0.017**
	(0.025)	(0.008)	(0.009)	(0.019)	(0.011)	(0.003)
Post * Distance	0.030	0.004	-0.019	0.043*	-0.015	-0.009**
	(0.027)	(0.009)	(0.011)	(0.021)	(0.012)	(0.003)
% Male Patients	0.154	0.028	0.047	0.069	0.074	0.002
	(0.092)	(0.035)	(0.039)	(0.076)	(0.052)	(0.013)
% Patients Age 85+	-0.072	0.027	-0.116**	0.020	-0.090**	-0.020*
	(0.064)	(0.025)	(0.032)	(0.052)	(0.034)	(0.009)
% Black Patients	0.094	-0.051	0.154**	-0.003	0.104	0.021
	(0.100)	(0.049)	(0.051)	(0.081)	(0.053)	(0.014)
% Hispanic Patients	0.314*	-0.025	0.077	0.248*	0.052	0.003
	(0.151)	(0.045)	(0.055)	(0.116)	(0.060)	(0.014)
% Paid by Medi-Cal	0.067	-0.069	-0.039	0.206	-0.108	-0.040*
	(0.146)	(0.040)	(0.047)	(0.134)	(0.076)	(0.016)
% Paid by Medicare	0.010	0.014	-0.119*	0.142	-0.104	-0.027
	(0.170)	(0.051)	(0.056)	(0.140)	(0.086)	(0.017)
Constant	2.727**	0.438**	0.491**	1.712**	0.929**	0.339**
	(0.181)	(0.055)	(0.110)	(0.149)	(0.138)	(0.036)
Observations	4896	4896	4896	4896	4896	4896
Number of SNFs	612	612	612	612	612	612
R-squared	0.337	0.259	0.256	0.333	0.293	0.161

Notes: * Significant at the 5% level. ** Significant at the 1% level. SNF clustered robust standard errors reported in parentheses. SNF fixed effects and fixed trends included. Coefficient estimates on selected explanatory variables not reported for brevity. RN = registered nurse, LVN = licensed vocational nurse, CNA = certified nurse aide, and HPRD2000 = hours per resident day under the 2000 calculation. Post = 1 if t=2000-2002 and = 0 for t<2000.

Table 2.4: Nurse Staffing Results for High Staffing SNFs

Variables	HPRD2000	RN HPRD	LVN HPRD	CNA HPRD	Licensed HPRD	% Licensed
Post	0.030	-0.056**	0.046*	0.042	-0.010	-0.003
	(0.065)	(0.021)	(0.019)	(0.046)	(0.025)	(0.004)
% Male Patients	0.348	0.131	0.050	0.161	0.182	0.002
	(0.286)	(0.099)	(0.070)	(0.188)	(0.109)	(0.016)
% Patients Age 85+	-0.086	0.010	-0.039	-0.051	-0.029	0.010
	(0.193)	(0.076)	(0.062)	(0.132)	(0.092)	(0.018)
% Black Patients	0.904	0.147	0.083	0.711*	0.231	-0.035
	(0.517)	(0.169)	(0.184)	(0.310)	(0.278)	(0.027)
% Hispanic Patients	0.488	0.401	-0.077	0.161	0.324	0.007
	(0.645)	(0.222)	(0.208)	(0.448)	(0.276)	(0.043)
% Paid by Medi-Cal	-0.347	-0.037	-0.031	-0.280*	-0.069	0.009
	(0.217)	(0.075)	(0.072)	(0.135)	(0.105)	(0.013)
% Paid by Medicare	0.045	0.025	0.104	-0.082	0.129	0.024
	(0.213)	(0.083)	(0.065)	(0.153)	(0.096)	(0.014)
Constant	5.092**	0.642**	0.992**	3.277**	1.634**	0.345**
	(0.538)	(0.151)	(0.184)	(0.361)	(0.272)	(0.054)
Observations	1600	1600	1600	1600	1600	1600
Number of SNFs	200	200	200	200	200	200
R-squared	0.178	0.171	0.211	0.160	0.199	0.169

Notes: * Significant at the 5% level. ** Significant at the 1% level. SNF clustered robust standard errors reported in parentheses. SNF fixed effects and fixed trends included. Coefficient estimates on selected explanatory variables not reported for brevity. RN = registered nurse, LVN = licensed vocational nurse, CNA = certified nurse aide, and HPRD2000 = hours per resident day under the 2000 calculation. Post = 1 if t=2000-2002 and = 0 for t<2000. High Staffing SNFs are facilities with HPRD2000 values greater than or equal to 3.2 in 1999.

Table 2.5: Effect of Staffing on the # of Patients Discharged Due to Death for Low Staffing SNFs

Variables	OLS	IV
HPRD2000	0.008	-6.137**
	(0.646)	(2.341)
% Male Patients	9.908*	11.087**
	(4.073)	(4.047)
% Patients Age 85+	3.471	2.362
	(3.823)	(3.511)
% Black Patients	6.854	7.642
	(5.348)	(5.015)
% Hispanic Patients	-6.919	-3.865
	(6.276)	(6.073)
% Paid by Medi-Cal	-3.474	-2.912
	(5.034)	(5.065)
% Paid by Medicare	5.895	6.248
	(6.350)	(6.219)
# of Licensed Beds	0.103	0.098
	(0.069)	(0.067)
Occupancy Rate	-0.013	-0.017
	(0.009)	(0.009)
Constant	20.227*	
	(8.646)	
Observations	4896	4896
Number of SNFs	612	612
First Stage F		170.0

Notes: * Significant at the 5% level. ** Significant at the 1% level. SNF clustered robust standard errors reported in parentheses. SNF fixed effects and fixed trends included. Coefficient estimates on county level unemployment rate, poverty rate, and linear time trend are not reported. Post = 1 if t=2000-2002 and = 0 for t<2000. HPRD2000 = hours per resident day under the 2000 calculation. Low Staffing SNFs are facilities with HPRD2000 values less than 3.2 in 1999.

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3 Child Support Enforcement and the Incidence of Single Motherhood

This paper examines how child support enforcement (CSE) reform created by the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 affects the incidence of single motherhood in the United States. Results demonstrate that the effect of CSE reform differs for women by educational attainment. Estimates reveal that the fraction of women who are single mothers would be 12.3 and 6.3 percent lower in the absence of CSE reform for women with less than 12 years of education and women with 12 years of education, respectively. Rises in single motherhood come from increases in marital dissolutions and decreases in marriage among women with out-of-wedlock births. These results suggest that CSE has an impact on household composition and the circumstances in which children are raised.

3.1 Introduction

This study examines how child support enforcement (CSE) affects the incidence of single motherhood in the United States. This paper provides implications on how recent CSE reform affects the types of households in which children are raised. I estimate the impact of CSE on subgroups of women by educational attainment to determine if certain segments of the female population are more or less likely to become single mothers.

My study improves upon previous research by using variation in when states implement CSE reform to estimate the effect of CSE on the incidence of single motherhood. Studies investigating the effect of CSE on marriage use data on state level characteristics from the Office of Child Support Enforcement (OCSE) including child support collection rates, paternity establishment rates, and number of child support orders (Heim, 2003; Nixon, 1997). Garfinkel et al. (2003) use similar measures to determine that CSE decreases non-marital birth rates. While these measures provide valuable information on how well a state collects child support, using this data to estimate the impact of CSE on marriage and fertility may be subject to reverse causality if rates of fertility or marriage directly affect child support collection. For instance, states with a high fraction of never married mothers could have low collection rates simply because these states have cases that require additional steps to collect child support. In particular, child support cases for never married custodial mothers require paternity establishment. In this paper, I explicitly demonstrate that using state level variation in the implementation of CSE to identify the impact of CSE on the likelihood of being a single mother is not subject to reverse causality.

Using March Current Population Survey (CPS) data for years 1992-2004, I estimate how CSE reform created by the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996 affects the probability of being a single mother for women age 20-45. I study the effects of these reforms on the incidence of single motherhood separately for women with less than a high school degree, high school graduates, and women with some college education or more. A priori, it is unclear how CSE reform will affect the likelihood of being a single mother. If CSE increases the likelihood that fathers will have to provide financial support to their current or potential children, then improvements in CSE may prevent married fathers from filing for divorce, deter men from having out-of-wedlock children, or even induce men to marry after having non-marital births. In contrast, better CSE might provide women with a stronger safety net and cause them to be more likely to pursue a divorce, have out-of-wedlock births, or forgo marriage after a non-marital birth.

Results from this study demonstrate that CSE reform causes the incidence of single motherhood to increase for women with less than a high school degree and female high school graduates.[1] I also find that CSE does not have a statistically significant effect on women with some college education. Although previous research determines that pre-PRWORA CSE decreases the incidence of out-of-wedlock births (Huang, 2002; Garfinkel et al., 2003), I determine that recent CSE reform deters marriage among low educated women with non-marital births and increases marital dissolutions of female high school graduates. These results show that CSE reform causes the likelihood of being raised in a single parent household to increase for children of low educated women.

The remainder of this paper will be organized as follows. Section 2 discusses previous literature, Section 3 gives background on CSE policy, Section 4 describes the data, Section 5 contains the empirical estimation and results, and Section 6 provides discussion.

3.2 Previous Literature

Previous literature provides conflicting results on the effect of CSE on divorce and non-marital birth rates. These studies generally use data prior to PRWORA and measure CSE using state level data on child support outcomes. Using state level child support outcomes to proxy for CSE is problematic if marriage and fertility directly affect child support outcomes. Some studies provide evidence that improvements in CSE decrease the incidence of non-marital births (Huang, 2002; Garfinkel et al., 2003) and divorce (Nixon, 1997), suggesting that CSE decreases rates of single motherhood. Nixon (1997) finds a small negative effect from stronger CSE on the likelihood of currently being divorced conditional on being married five years prior and having at least one child under the age of 18. In contrast, Heim (2003) uses state level vital statistics data to determine that CSE does not have a statistically significant effect on divorce rates. Other studies find that CSE increases rates of single motherhood by decreasing remarriage rates

[1]Female high school graduates include women with 12 years of education only and exclude women with some college attendance.

of divorced women (Folk et al., 1992) and decreasing marriage among couples who have non-marital births (Carlson et al., 2005). In particular, Folk et al. (1992) determine that women who are more likely to receive child support payments and women who are more likely to receive large child support payments are less likely to remarry if they have not remarried within five years of a divorce. A study by Carlson et al. (2005) provides some evidence that CSE in the late 1990s deters couples who have non-marital births from marrying in the future.

My study improves upon previous research by using variation in when states implement CSE reform to estimate the effect of CSE on the incidence of single motherhood. By using variation in timing, I bypass problems of reverse causality. In addition, I contribute to the existing literature by estimating the effects of CSE separately for women by education, which allows me to determine whether subgroups of women are affected differently, and if so, what the implications may be for children growing up in these households.

3.3 Child Support Enforcement Policy

In 1974, federal and state CSE offices were created to collect child support on behalf of custodial parents on welfare. To receive welfare benefits, participants were required to comply with CSE agencies. Although the federal Office of Child Support Enforcement (CSE) states their goal as securing the "well-being of children by assuring that assistance in obtaining support...is available to children," federal intervention in child support collection originated as a way to make sure welfare recipients were not funded twice through welfare and child support. Child support collection functioned as a revenue generating process in which state governments retained child support collected on behalf of welfare participants. As a result, non-custodial fathers faced disincentives to pay child support to mothers on welfare because their payments went directly to the government. The 1984 CSE amendments extended services of CSE agencies to non-welfare participants. In recognition of the disincentives associated with child support payments in welfare cases, the Deficit Reduction Act of 1984 amended state laws to allow the first $50 of monthly child support payments collected on behalf of welfare participants to go directly to the custodial parent. This $50 is commonly referred to as the child support pass through.

The Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996 created extensive reform in CSE by requiring states to improve paternity establishment rates, impose better technology to locate parents, streamline the imposition of penalties, and create new penalties. States were also granted the option to maintain or abolish their child support pass through policies. Noncompliance with PRWORA reform did not result in direct fines; however, states had financial incentives to comply since federal funding for state welfare programs was directly tied with state level child support performance.

I use Office of Child Support Enforcement State Plans to obtain data on when PRWORA CSE reforms were instituted in each state. While the provisions of PRWORA apply to each state, the dates in which these laws passed vary across states. To measure CSE reform, I construct a variable called Fraction of Laws Passed which equals the number of laws passed in a state divided by 11, the total number of possible laws.[2]

I demonstrate that Fraction of Laws Passed is not subject to reverse causality by collapsing pre-reform data for years 1994-1995 to the state level and using a multinomial logistic estimation to test whether pre-reform state level characteristics affect when states implement reform. There are three main waves in which states institute reforms occurring in 1997, 1998, and 1999. I categorize states as 1997 movers, 1998 movers, and 1999 movers based on the year when a state passes the most laws. There are 8 1997 movers, 33 1998 movers, and 9 1999 movers. States generally pass a majority of laws, over 80%, when they move. Therefore, the main source of variation in Fraction of Laws Passed comes from when reforms are implemented, particularly when the majority of laws are passed. Results in Table 1.8 reveal that states with a higher fraction of women and lower fraction

[2]Although I have data on 16 CSE laws, these measures are calculated using a subset of 11 laws. See the Appendix in Chapter 1 for detailed descriptions of these laws. These 11 laws are policies in which the majority, if not all, are passed when a state moves. The remaining 5 laws are passed sporadically across and within each state. Including these additional 5 laws in the estimation yields similar, but less precise results.

of Republicans in the state legislature are more likely to be a 1997 mover relative to a 1999 mover. In addition, states with a higher fraction of women in the state legislature are more likely to be a 1998 mover relative to a 1999 mover. Having a higher fraction of women in a state legislature could cause states to implement reform sooner because women might be more sympathetic to the single mother population as possible mothers themselves. In general, Democrats are more likely to support increases in expenditures on social welfare programs than Republicans, which explains why states with a higher fraction of Democrats in the state legislature may be quicker to implement new reforms. Pre-reform state level child support outcomes, as measured by the fraction of child support cases with collections, does not significantly affect when states pass CSE laws. Furthermore, pre-reform state level female demographic variables have a jointly insignificant effect on state mover status. Thus, the implementation of laws is not found to be significantly correlated with pre-reform child support outcomes or the composition of women, and there is no evidence suggesting that the variation in Fraction of Laws Passed comes from changes in fertility, marriage, or child support trends.

3.4 Data

I use annual March CPS data for years 1992-2004. The sample consists of women age 20-45. Because this study examines the effects of child support reform separately by years of education, I exclude teenagers from the sample because these women are unlikely to have had the opportunity to graduate high school or attend college. A woman is a single mother if she is currently unmarried and has a child young enough for the mother to be affected by child support reform. As discussed in Section 3.3, states generally began enacting child support reform in 1997. Women whose youngest child is aged 14 or older in 1992 would not be affected by CSE reform because their children will be older than 18 in 1997, causing them to be ineligible for child support. By a similar reasoning, a woman's youngest child must be under age 15 in 1993, 16 in 1994, 17 in 1995, and 18 in 1996 to have an age eligible child, where an age eligible child refers to having a child young enough for the mother to be affected by child support reform. Based on this definition, roughly 14% of women in the sample are single mothers, and the frequency of single motherhood decreases with years of education as reported in Table 3.1. The percentage of women with less than a high school degree who are single mothers is roughly 30%, over 2.5 times the percentage for women with some college education. Comparing demographic characteristics by educational attainment also demonstrates that lower educated women are more likely to be non-white, live in a city, and have an age eligible child.

Table 3.2 reports the fraction of women who are single mothers by pre (1992-1996) and post (2000-2004) reform time periods. The post reform percentage of women who are single mothers is larger than the pre reform average. This appears to be a result of a rise in the fraction of never married single mothers. As expected, women with less than a high school degree are more likely to be never married single mothers than high educated women both pre and post reform. Divorce rates are roughly equal for women with less than a high school degree and female high school graduates after CSE reform occurs, but are still much higher than the divorce rate for women with some college education. In the next section, I use empirical methods to determine how much of these changes in marriage and fertility are attributed to CSE reform.

3.5 Empirical Estimation

To examine the effect of CSE on the incidence of single motherhood, I estimate the following probit model:

$$\Pr(\text{Single Mother})_{ist} = \Phi(XB) \tag{3.5.1}$$

where

$$XB = A_s + B_t + A_s t + \beta \text{Fraction of Laws Passed}_{st} + \Gamma X_{ist} + \Pi S_{st} + \epsilon_{ist} \tag{3.5.2}$$

X_{ist} is a vector of individual level time-varying characteristics including age, race, and city residence status for individual i in state s in year t. S_{st} is a vector of state level time-varying characteristics consisting of welfare reform measures, the maximum earned income tax credit for a mother with two children, the unemployment rate, average female hourly wage, and average male hourly wage. I include state fixed effects, A_s, to account for time-invariant state level characteristics and year fixed effects, B_t, to account for time-varying characteristics shared across states. $A_s t$ are state fixed trends which account for state specific variables trending linearly during this time period. The dependent variable of interest is the Fraction of Laws Passed and its coefficient β measures the effect of instituting all of the CSE laws.

3.5.1 Results

Results in Table 3.3 demonstrate that Fraction of Laws Passed has a statistically significant and positive effect on the incidence of single motherhood for women with less than a high school degree and female high school graduates. In addition, CSE has no significant effect on rates of single motherhood for women with at least some college education. Passing all reform measures causes the likelihood of being a single mother to increase by 4.9 percentage points for women with less than a high school degree, a result that is statistically significant at the 5% level. For female high school graduates, instituting CSE reform causes the likelihood of being a single mother to increase by 2.1 percentage points, an estimate that is statistically significant at the 1% level.

To better understand the magnitude of the effect of CSE reform on the incidence of single motherhood, I compare predicted probabilities of being a single mother to observed probabilities in Table 3.4. I calculate the predicted probability of being a single mother in the absence of CSE reform by setting Fraction of Laws Passed equal to zero during the post reform time period 2000-2004. Without CSE reform, the likelihood of being a single mother is 3.6 percentage points, or 12.3 percent, lower than the observed value for women with less than a high school degree. In the absence of CSE reform, the probability that a female high school graduate is a single mother is 1.3 percentage points, or 6.3 percent, lower than the observed probability post reform.

To determine the mechanism by which CSE affects single motherhood, I examine how CSE impacts fertility and marital status of low educated women. Unreported results show that CSE does not have a significant effect on the likelihood of having an age eligible child for women with less than 12 years of education or women with 12 years of education.[3] Therefore, the promotion of single motherhood is not caused by changes in fertility. Because CSE does not significantly affect fertility, I use a multinomial logistic model to estimate the impact of CSE on the likelihood of being married, divorced, and never married conditional on having an age eligible child.

The coefficient estimates reported in Table 3.5 are relative risk ratios with being married as the base outcome. The coefficient estimates for Fraction of Laws Passed are all greater than 1, meaning that there is evidence that CSE reform increases the relative risks of being both divorced and never married relative to being married conditional on having an age eligible child. For women with less than a high school degree, CSE causes the relative risk of being never married relative to being married to increase by 1.43. To interpret the magnitude of this increase in never married single motherhood, I estimate the predicted probabilities of being married, divorced, and never married conditional on having children in Table 3.6. In the absence of CSE reform, the percentage of mothers with less than 12 years of education who are never married would be 3.7 percentage points lower or 14% less than the observed average. Although the coefficient estimate on Fraction of Laws Passed is greater than 1 when estimating its impact on the likelihood that a woman with less than 12 years of education is divorced relative to being married, the standard errors are too large to make statistical inferences.

Results show that CSE reform increases the relative risk of being divorced relative to being married by 1.19 for mothers with 12 years of education. Without CSE, the percentage of mothers with 12 years of education who are divorced would be 2.2 percentage points lower, or 14% lower than the observed percentage. The

[3]Fraction of Laws Passed also does not have a significant effect on fertility of women with some college education.

coefficient estimate on Fraction of Laws Passed is greater than 1 when estimating the impact of CSE reform on being never married relative to married, suggesting that CSE reform also increases non-marital births. However, the standard errors are too large to make statistical inferences. These results demonstrate that CSE reform promotes divorce among women with 12 years of education. For mothers with less than 12 years of education, CSE deters marriage among mothers with non-marital births.

To further investigate whether CSE decreases marriage among women with non-marital births, I estimate the effect of CSE on the marital status of women who have children aged 3 or under.[4] By estimating the effect of CSE on marital status conditional on having young children, I isolate the effect of child support reform on marriage decisions of women with recent births. Results in Table 3.7 show that CSE does not significantly affect the marital status of mothers of young children who have less than 12 years of education. This result does not necessarily contradict previous estimates because the sample size of mothers with young children who also have less than 12 years of education could be too small to obtain precise estimates. In fact, comparing predicted probabilities with observed values in Table 3.8 reveals that the incidence of being never married and divorced conditional on having a child aged 0 to 3 would be lower without CSE reform, which supports the previous analysis on marital history.

Results also demonstrate that women with 12 years of education are more likely to be never married relative to being married conditional on having a young child. Thus, CSE reform also reduces the likelihood of marriage for high school graduates with recent out-of-wedlock births. The relative risk of being a never married single mother is 1.53 times more likely than being a married mother for mothers with young children and 12 years of education. Without CSE reform, the percentage of female high school graduates with children aged 0 to 3 who are never marred would be 4.6 percentage points or almost 20% lower than the observed percentage. Since CPS data is not a longitudinal dataset, I am unable to track women over time to determine who eventually marries and who remains

[4]Fraction of Laws Passed does not affect the likelihood of having a child aged 3 or under for subgroups of women by years of education.

unmarried after a non-marital birth. Even though I cannot make inferences on the flow of marriage rates, I determine that CSE reform increases the stock of never married single mothers.

3.5.2 Robustness Checks

To make sure that women do not select into different educational attainment groups, I estimate a multinomial logistic equation to determine whether CSE reform affects the likelihood of having less than 12 years of education, 12 years of education, or more than 12 years of education. Fraction of Laws Passed does not have a significant effect on educational attainment of women, demonstrating that CSE reform does not cause women to select into different education groups.

To show that results discussed in Section 3.5.1 are not caused by secular changes in marriage or fertility trends, I estimate the effect of CSE reform on marriage and fertility for subgroups of women who are predicted to be unaffected by changes in child support policy. First, I estimate the effect of CSE reform on the marital status of childless women. Presumably, child support should not affect the likelihood of marriage or divorce for women who are ineligible to collect child support. Second, I estimate the effect of CSE reform on fertility of married women. Again, since married women are ineligible to collect child support, changes in CSE should not affect the decision to have a child for this group. Unreported results demonstrate that CSE does not affect marriage or fertility of women in these respective samples. Additional estimations using subgroups of women by educational attainment also yield statistically insignificant results. Therefore, I confirm that the main empirical results are not caused by overall changes in fertility or marriage trends.

3.6 Discussion

This study demonstrates that recent CSE reform increases the incidence of single motherhood for women with a high school degree or less. In the absence of CSE reform, rates of single motherhood would be 12.3% and 6.3% lower for women with less than 12 years of education and women with 12 years of education, respectively. Results indicate that the increase in the probability of single motherhood for women with less than a high school degree stems from a rise in never married single motherhood. The increase in the probability of single motherhood for female high school graduates comes from a promotion of marital dissolutions and never married single motherhood.

At face value, these results seem to suggest that CSE policy causes low educated women to be worse off by promoting the incidence of single motherhood. Although the impact on child well-being depends on each family's circumstance before and after a divorce, studies typically find that single mother households are worse off financially after a marital dissolution (Bartfeld, 2000; Weiss and Willis, 1993; Weiss, 1984). Furthermore, children of low educated women are increasingly likely to be raised in single parent households, implying that CSE could increase disparities between children with lower and higher educated mothers. However, to determine how these changes in marriage decisions affect well-being of low educated single mothers, research needs to identify which sub sample of low educated women are being induced to become single mothers and who these women are deciding not to marry. If these women are choosing to forgo marriage with volatile or abusive men, then these single mother families might be better off as a result of CSE reform. Future research should assess how CSE reform affects marriage matches among low educated women and how these changes in household formation impact well-being.

Table 3.1: Summary Statistics By Years of Education

Variable	All Women		<12 years		=12 years		>12 years	
	Mean	(Std. Dev.)	Mean	(Std. Dev.)	Mean	(Std. Dev.)	Mean	(Std. Dev.)
Age	33.158	(7.361)	32.469	(7.57)	33.232	(7.405)	32.765	(7.390)
Live in city	0.238	(0.426)	0.317	(0.465)	0.224	(0.417)	0.25	(0.433)
White	0.721	(0.448)	0.525	(0.499)	0.717	(0.45)	0.773	(0.419)
Hispanic	0.117	(0.321)	0.23	(0.421)	0.087	(0.282)	0.062	(0.241)
Black	0.117	(0.321)	0.21	(0.407)	0.167	(0.373)	0.122	(0.328)
Less than H.S.	0.086	(0.281)	1	(0)	0	(0)	0	(0)
High school graduate	0.337	(0.473)	0	(0)	1	(0)	0	(0)
Some college or more	0.577	(0.494)	0	(0)	0	(0)	1	(0)
Any Children	0.6	(0.49)	0.66	(0.474)	0.629	(0.483)	0.52	(0.5)
Married	0.586	(0.493)	0.484	(0.5)	0.586	(0.493)	0.567	(0.496)
Never married	0.275	(0.447)	0.315	(0.465)	0.251	(0.434)	0.315	(0.465)
Divorced or Separated	0.139	(0.346)	0.201	(0.401)	0.163	(0.369)	0.118	(0.323)
Single Mother	0.155	(0.362)	0.293	(0.455)	0.189	(0.391)	0.112	(0.316)
Observations	337270		29166		113664		194440	

Notes: Summary statistics calculated using March Current Population Survey data. Means are weighted by person weights.

Table 3.2: Pre Post Means by Years of Education

	1992-1996		2000-2004	
All				
Single Mother	0.144	(0.351)	0.158	(0.364)
Single Mother, Never Married	0.057	(0.233)	0.075	(0.264)
Single Mother, Divorced	0.087	(0.281)	0.082	(0.275)
Observations	130686		142840	
< 12 years				
Single Mother	0.278	(0.448)	0.293	(0.455)
Single Mother, Never Married	0.141	(0.348)	0.175	(0.38)
Single Mother, Divorced	0.137	(0.344)	0.118	(0.323)
Observations	13426		10567	
= 12 years				
Single Mother	0.171	(0.376)	0.206	(0.405)
Single Mother, Never Married	0.073	(0.259)	0.105	(0.307)
Single Mother, Divorced	0.098	(0.297)	0.101	(0.302)
Observations	46555		45448	
> 12 years				
Single Mother	0.103	(0.304)	0.116	(0.321)
Single Mother, Never Married	0.033	(0.177)	0.048	(0.213)
Single Mother, Divorced	0.07	(0.256)	0.069	(0.253)
Observations	70705		86825	

Notes: Summary statistics calculated using March Current Population Survey data. Means are weighted by person weights.

Table 3.3: Effect of Child Support Enforcement on Pr(Single Mother by Years of Education)

Variable	All	<12 years	=12 years	>12 years
Fraction of Laws Passed	0.0091*	0.0490*	0.0205**	-0.0013
	(0.0040)	(0.0227)	(0.0073)	(0.0044)
TANF Implemented	-0.0015	-0.0095	-0.0050	0.0015
	(0.0046)	(0.0256)	(0.0095)	(0.0042)
LN(Welfare benefit)	-0.0144	-0.1102	-0.0164	0.0053
	(0.0235)	(0.1119)	(0.0314)	(0.0204)
TANF Time Limit	-0.0026	-0.0102	-0.0049	-0.0013
	(0.0033)	(0.0162)	(0.0063)	(0.0036)
Child Support Pass Through	-0.0046	-0.0130	-0.0076	-0.0023
	(0.0038)	(0.0234)	(0.0073)	(0.0044)
Maximum EITC (in 100s)	0.0004	0.0043	-0.0015	0.0005
	(0.0010)	(0.0028)	(0.0024)	(0.0009)
State Unemployment Rate	0.0042**	-0.0045	0.0069**	0.0035
	(0.0015)	(0.0057)	(0.0026)	(0.0021)
Avg Female Wage (in 100s)	-0.0036	0.0999**	-0.0039	-0.0116*
	(0.0025)	(0.0273)	(0.0073)	(0.0053)
Avg Male Wage (in 100s)	-0.0322	-0.0498	-0.0173	-0.0431**
	(0.0184)	(0.0823)	(0.0350)	(0.0164)
Age	0.0317**	0.0437**	0.0323**	0.0271**
	(0.0008)	(0.0039)	(0.0019)	(0.0012)
Hispanic	0.0642**	0.0756**	0.0781**	0.0770**
	(0.0146)	(0.0256)	(0.0164)	(0.0131)
Black	0.2248**	0.2923**	0.2611**	0.2124**
	(0.0055)	(0.0130)	(0.0073)	(0.0072)
Log Likelihood	-132282	-15950.78	-51221.03	-65649.06
Pseudo R-squared	0.089	0.087	0.067	0.055
N	337270	29166	113664	194440

Notes: Estimated using a logistic regression. Coefficient estimates reported as marginal effects. Selected estimates omitted for brevity. State clustered robust standard errors reported in parentheses. * significant at the 5% level and ** significant at the 1% level. TANF stands for Temporary Assistance for Needy Families and EITC stands for Earned Income Tax Credit.

Table 3.4: Predicted Probabilities of Being a Single Mother

	1992-1996		2000-2004	
All				
Single Mother, Observed	0.144	(0.351)	0.158	(0.364)
Single Mother, Predicted	0.141	(0.106)	0.164	(0.113)
Single Mother, Without CSE	0.141	(0.106)	0.156	(0.104)
Observations	130686		142840	
$<$ **12 years**				
Single Mother, Observed	0.278	(0.448)	0.293	(0.455)
Single Mother, Predicted	0.273	(0.147)	0.30	(0.158)
Single Mother, Without CSE	0.273	(0.147)	0.257	(0.15)
Observations	13426		10567	
$=$ **12 years**				
Single Mother, Observed	0.171	(0.376)	0.206	(0.405)
Single Mother, Predicted	0.167	(0.102)	0.213	(0.116)
Single Mother, Without CSE	0.166	(0.102)	0.193	(0.11)
Observations	46555		45448	
$>$ **12 years**				
Single Mother, Observed	0.103	(0.304)	0.116	(0.321)
Single Mother, Predicted	0.101	(0.069)	0.123	(0.075)
Single Mother, Without CSE	0.102	(0.069)	0.124	(0.076)
Observations	70705		86825	

Notes: Predicted probabilities calculated using coefficient estimates from the logistic regression in Table 3.3. Predicted probabilities without CSE calculated setting Fraction of Laws Passed equal to zero. Means are weighted by person weights.

Table 3.5: Multinomial Logistic results on Marital Status Conditional on Having an Age Eligible Child by Years of Education

	<12 years		=12 years	
Variable	Divorced	Never Married	Divorced	Never Married
Fraction of Laws Passed	1.2583	1.4292*	1.1897*	1.1390
	(0.1847)	(0.2321)	(0.0881)	(0.1362)
TANF Implemented	0.9482	0.9111	0.9006	1.1170
	(0.1549)	(0.1394)	(0.0805)	(0.1316)
LN(Welfare benefit)	0.9968	0.3014	0.6466	1.0462
	(0.5249)	(0.2609)	(0.2174)	(0.3397)
TANF Time Limit	0.8959	0.9178	0.8986	1.0311
	(0.1245)	(0.0875)	(0.0491)	(0.0818)
Child Support Pass Through	0.8363	1.1124	0.8449*	1.0999
	(0.1265)	(0.1517)	(0.0606)	(0.0995)
Maximum EITC (in 100s)	1.0635*	1.0036	0.9863	0.9879
	(0.0290)	(0.0216)	(0.0170)	(0.0172)
State Unemployment Rate	0.9927	1.0133	1.0617	1.0695
	(0.0502)	(0.0506)	(0.0342)	(0.0414)
Avg Female Wage (in 100s)	2.0788**	2.2642**	1.0397	0.6809
	(0.5128)	(0.2544)	(0.0742)	(0.1803)
Avg Male Wage (in 100s)	0.7218	1.5119	1.0080	0.7067
	(0.3627)	(1.0291)	(0.2569)	(0.3184)
Age	1.1683**	0.7674**	1.0774**	0.6541**
	(0.0467)	(0.0301)	(0.0227)	(0.0190)
Hispanic	0.9226	1.6014**	1.2420*	2.2055**
	(0.1467)	(0.2198)	(0.1250)	(0.2484)
Black	2.8223**	15.9804**	2.8001**	15.5669**
	(0.2405)	(2.4319)	(0.1562)	(0.8640)
Constant	0.0061	9.03e+04	0.7993	359.3459*
	(0.0220)	(5.46e+05)	(1.7868)	(842.9741)
Log Likelihood	-16121.91		-50733.81	
Pseudo R-squared	0.160		0.143	
N	19706		74062	

Notes: Estimated with being married as the base outcome. Coefficient estimates reported as relative risk ratios. Selected estimates omitted for brevity. State clustered robust standard errors reported in parentheses. * significant at the 5% level and ** significant at the 1% level. TANF stands for Temporary Assistance for Needy Families and EITC stands for Earned Income Tax Credit.

Table 3.6: Predicted Probabilities of Being Married, Divorce, and Never Married Conditional on Having an Age Eligible Child

	Type	1992-1996		2000-2004	
< 12 years					
Married	Observed	0.574	(0.495)	0.549	(0.498)
Married	Predicted	0.579	(0.209)	0.556	(0.223)
Married	Without CSE	0.579	(0.209)	0.610	(0.224)
Divorced	Observed	0.210	(0.407)	0.181	(0.385)
Divorced	Predicted	0.206	(0.082)	0.175	(0.076)
Divorced	Without CSE	0.206	(0.082)	0.158	(0.069)
Never Married	Observed	0.217	(0.412)	0.270	(0.444)
Never Married	Predicted	0.216	(0.223)	0.269	(0.250)
Never Married	Without CSE	0.215	(0.222)	0.233	(0.240)
Observations		8830		7267	
= 12 years					
Married	Observed	0.720	(0.450)	0.676	(0.468)
Married	Predicted	0.725	(0.171)	0.679	(0.194)
Married	Without CSE	0.725	(0.171)	0.705	(0.193)
Divorced	Observed	0.161	(0.368)	0.160	(0.366)
Divorced	Predicted	0.158	(0.050)	0.156	(0.048)
Divorced	Without CSE	0.158	(0.050)	0.138	(0.044)
Never Married	Observed	0.120	(0.324)	0.165	(0.371)
Never Married	Predicted	0.118	(0.167)	0.165	(0.202)
Never Married	Without CSE	0.118	(0.167)	0.157	(0.197)
Observations		28503		31198	

Notes: Predicted probabilities calculated using coefficient estimates from the multinomial logistic regression in Table 3.5. Predicted probabilities without CSE calculated setting Fraction of Laws Passed equal to zero. Means are weighted by person weights.

Table 3.7: Multinomial Logistic results on Marital Status Conditional on Having Children Aged 0-3 by Years of Education

	<12 years		=12 years	
Variable	Divorced	Never Married	Divorced	Never Married
Fraction of Laws Passed	1.6549	1.5868	1.0807	1.5255*
	(0.4459)	(0.4052)	(0.1891)	(0.2929)
TANF Implemented	0.8771	0.9452	0.8099	1.2518
	(0.2180)	(0.1971)	(0.1295)	(0.2052)
LN(Welfare benefit)	0.2862	0.1376	0.2963**	0.5083
	(0.3722)	(0.1694)	(0.1395)	(0.2344)
TANF Time Limit	0.9476	0.9448	0.9171	1.0225
	(0.2338)	(0.2188)	(0.1471)	(0.2037)
Child Support Pass Through	0.9494	1.1050	0.9674	1.2336
	(0.2115)	(0.3011)	(0.1694)	(0.1629)
Maximum EITC (in 100s)	1.1263*	1.0073	1.0011	0.9768
	(0.0587)	(0.0344)	(0.0346)	(0.0176)
State Unemployment Rate	0.8806	1.0758	1.0769	1.0475
	(0.0809)	(0.0679)	(0.0737)	(0.0559)
Avg Female Wage (in 100s)	1.0793	1.8049*	1.2036	0.7118
	(0.9614)	(0.5065)	(0.2795)	(0.1954)
Avg Male Wage (in 100s)	1.2715	1.6117	1.2690	0.6382
	(1.6777)	(1.6263)	(1.0348)	(0.3233)
Age	1.0214	0.6824**	0.8647**	0.4969**
	(0.0792)	(0.0261)	(0.0372)	(0.0240)
Hispanic	0.9910	1.3960*	1.3025	1.8163**
	(0.2226)	(0.1993)	(0.1930)	(0.2092)
Black	3.8369**	17.3462**	3.0831**	13.4522**
	(0.5909)	(3.4729)	(0.2687)	(0.9836)
Log Likelihood	-5896.226		-15691.53	
Pseudo R-squared	0.182		0.170	
N	7509		24486	

Notes: Estimated with being married as the base outcome. Coefficient estimates reported as relative risk ratios. Selected estimates omitted for brevity. State clustered robust standard errors reported in parentheses. * significant at the 5% level and ** significant at the 1% level. TANF stands for Temporary Assistance for Needy Families and EITC stands for Earned Income Tax Credit.

Table 3.8: Predicted Probabilities of Being Married, Divorce, and Never Married Conditional on Having a Child Aged 0-3

	Type	1992-1996		2000-2004	
$<$ 12 years					
Married	Observed	0.560	(0.497)	0.512	(0.500)
Married	Predicted	0.564	(0.237)	0.518	(0.257)
Married	Without CSE	0.565	(0.236)	0.599	(0.267)
Divorced	Observed	0.165	(0.371)	0.106	(0.308)
Divorced	Predicted	0.163	(0.083)	0.102	(0.062)
Divorced	Without CSE	0.162	(0.082)	0.076	(0.048)
Never Married	Observed	0.276	(0.447)	0.382	(0.486)
Never Married	Predicted	0.274	(0.245)	0.380	(0.272)
Never Married	Without CSE	0.273	(0.244)	0.325	(0.271)
Observations		3717		2536	
$=$ 12 years					
Married	Observed	0.726	(0.446)	0.673	(0.469)
Married	Predicted	0.733	(0.198)	0.678	(0.229)
Married	Without CSE	0.733	(0.198)	0.722	(0.214)
Divorced	Observed	0.108	(0.310)	0.090	(0.286)
Divorced	Predicted	0.104	(0.038)	0.086	(0.030)
Divorced	Without CSE	0.104	(0.038)	0.087	(0.031)
Never Married	Observed	0.167	(0.373)	0.237	(0.425)
Never Married	Predicted	0.163	(0.193)	0.237	(0.234)
Never Married	Without CSE	0.162	(0.193)	0.191	(0.214)
Observations		10639		9463	

Notes: Predicted probabilities calculated using coefficient estimates from the multinomial logistic regression in Table 3.7. Predicted probabilities without CSE calculated setting Fraction of Laws Passed equal to zero. Means are weighted by person weights.

3.7 References

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