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WAGES, VIOLENCE AND HEALTH IN THE HOUSEHOLD

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ABSTRACT

Three quarters of all violence against women is perpetrated by domestic partners. I study both the economic causes and consequences of domestic violence. I find that decreases in the male-female wage gap reduce violence against women, consistent with a household bargaining model. The relationship between the wage gap and violence suggests that reductions in violence may provide an alternative explanation for the well-established finding that child health improves when mothers control a greater share of the household resources. Using instrumental variable and propsensity score techniques to control for selection into violent relationships, I find that violence against pregnant women negatively affects the health of their children at birth. This work sheds new light on the health production process as well as observed income gradients in health and suggests that in addition to addressing concerns of equity, pay parity can also improve the health of American women and children via reductions in violence.

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I. Introduction

Everyday 14 thousand women in the US are battered and four are killed by their intimate partners, with poor and disadvantaged women disproportionately affected. The estimated costs of domestic violence in terms of medical care and declines in productivity exceed $5.8 billion annually (CDC, 2003), prompting former Surgeon General C. Everett Koop to label domestic violence "the single most important health issue in the US." In this paper I examine some of the economic causes and consequence of domestic violence, analyzing both the impact of women's wages on domestic violence and the impact of domestic violence on child health. This work makes two contributions to the existing literature. First it establishes a negative causal relationship between women's wages and violence. Second, it identifies external costs associated with violence against women not previously considered and quantified: that the children of women who are the victims of violence suffer worse birth outcomes. Given the importance of birth outcomes in determining adult education and income (Black, Devereux and Salvanes, 2007), these results suggest that the higher levels of violence against poor women may contribute to the intergenerational transmission of economic status.

Violence and Women's Wages

Existing empirical research based on survey data has generally found that women with lower wages experience more violence. However, this work is limited in three respects. First, these studies fail to establish a causal relationship between domestic violence and women's wages by, for example, failing to account for the potential for omitted variable bias or reverse causality. Second, these studies focus largely on the woman's own wage when a household bargaining model suggests not only that women's <u>relative</u> wage matters but that <u>potential</u>, not actual, wages determine bargaining power and levels of violence. Third, they are based on data from household surveys which are prone to non-random underreporting.

To overcome these shortcomings, I employ two strategies. First, when using individual survey data, I instrument for women's income with characteristics of the local labor market conditions. To do so I take advantage of the fact that certain industries have traditionally been dominated by women (e.g., services) and others by men (e.g., construction). I create measures of local labor market conditions for men and women based on the industrial structure of the county and wage changes in the industries dominant in each county. Based on these measures, I find that as local labor market conditions for women improve, violence against them declines. I follow this with an analysis based on a new source of data on violence: administrative data on female hospitalization for assault. These data represent an improvement over individual level survey data because they do not rely on self-reports of violence and include the universe of all women in California (roughly 15 million individuals). I find that the improvement in local labor market conditions faced by women over the period 1990-2003 explains ten percent of the decline in violence against women witnessed over this period.

These findings are consistent with a simple model of household bargaining in which an increase in a woman's relative income increases her bargaining power and leads to a reduction in violence against her. The findings are inconsistent with models of "male backlash" developed by sociologists that predict that as women's wages increase, violence against them increases because men feel their traditional gender role threatened. They are also inconsistent with the model of exposure reduction developed by criminologists that predicts that as the labor force participation of women in creases, violence against them may decline because women spend less time with their violent partners. I find that the reductions in violence are most likely to occur during non-working hours, which is inconsistent with exposure reduction.

Violence and Child Health

Based on these findings I explore the impact of violence against women on child health at birth. Previous work based on developing countries has shown that child health improves when mothers control a greater share of the household resources (Thomas, 1990). This finding has been largely attributed to women's greater preferences for children and increased material investments in them, though there is little empirical support for this mechanism. The research presented here suggests another potential mechanism: reductions in violence. Violence during pregnancy can adversely affect the developing fetus via blunt trauma to the maternal abdomen (Silverman et al, 2006). Using a unique dataset that includes birth outcomes for all women in California and admissions to the hospital for an assault while pregnant, I explore the importance of violence as a mechanism behind the link between maternal resources and child well-being. Using instrumental variable and propensity score techniques to account for non-random selection into violent relationships, I find that reductions in violence against women result in improved birth outcomes. These results provide evidence of important externalities associated with domestic violence not previously quantified. In addition, given that poor women suffer more violence, this work sheds new light on observed income gradients in health.

The rest of this paper is laid out as follows: in section II I review the existing literature on domestic violence, discuss the shortcomings of existing empirical work and describe a bargaining model that incorporates violence, in section III I present the analysis of the impact of female income on domestic violence based on individual survey data, in section IV I present the analysis of the impact of women's relative labor market conditions on violence using administrative data on hospitalizations, section V contains the analysis of the impact of violence on birth outcomes, and section VI concludes.

II. Background on Domestic Violence

A. Prevalence of Domestic Violence and Risk Factors

Most estimates of domestic violence in the US come from the National Violence Against Women (NVAW) survey fielded in 1994. These data reveal an annual incidence of 2 percent, a lifetime incidence of 25 percent and that intimate partners are responsible for three fourths of all violence against women over the age of 18 (Tjaden and Thoennes, 1998). Disadvantaged women face much higher risks of abuse. Women with income below $10,000 annually report rates of domestic violence that are five times those with annual income greater than $30,000 (BJS, 1994). In addition, black women are at significantly greater risk of violence and conditional on violence are subject to more severe attacks (Rennison and Welchans, 2000). Also at greater risk are young women between the ages of 20 and 34.

The National Crime Victimization Survey (NCVS) is the only survey that has allowed tracking over time and findings suggests that the rates of domestic violence against women declined by 50 percent between 1993 and 2001.[1]

B. Theories of the Relationship Between Economic Status and Violence

Most of the existing research on domestic violence has been conducted by criminologists and sociologists. Criminologists have developed a theory of the relationship between employment and domestic violence referred to as exposure reduction. This theory posits that the increase in employment among either men or

[1] Intimate partner violence against men has also dropped over this period. Criminologists have suggested that the decline among men may be attributable to declines in the number of women in abusive relationships who kill their partners/abusers in defense (Dugan, Nagin and Rosenfeld, 1999). Due to a change in survey design, estimates prior to 1993 are not comparable to those obtained post 1993.

women will reduce domestic violence simply by reducing the time partners spend together, (Dugan, Nagin and Rosenfeld, 1999).

Two theories prominent in the sociological literature predict that as women's financial independence increases, violence against them should increase. The first theory is one of male backlash against increasing female independence associated with their increased employment and personal income. According to Macmillan and Gartner (1999), a wife's independence "signifies a challenge to culturally prescribed norm of male dominance and female dependence. Where a man lacks this sign of dominance, violence may be a means of reinstating his authority over his wife." The second theory derives from exchange theory and views domestic violence as one of the two sides of a reward/punishment approach to influence (Molm, 1989). Under this scenario, individuals possess two sources of power: transferring resources (rewards) and violence (punishment). As a husband's ability to influence his wife's behavior by transferring resources (rewards) diminishes when his income decreases relative to hers, he is more likely to rely on punishment which may include violence.

Theories of male backlash and exchange theory which predict that increases in women's wages lead to an increase in violence are problematic because they ignore the individually rational constraint faced by women in abusive relationships. That is, as their income increases, women are more likely to end the partnership if transfers decline and abuse continues or escalates.

Economic theories of household bargaining are consistent with the individually rational constraint but generally do not incorporate violence. Exceptions are Bloch and Rao (2002) and Farmer and Tiefenthaler (1997). Bloch and Rao (2002) incorporates asymmetric information and signaling in a model of noncooperative bargaining to explain why a woman from a wealthy family in India is subject to greater violence by her husband in an effort to extract more resources from her family. This model, however, does not fit the experience of women in the US. Farmer and Tiefenthaler (1997) present a particular case of a non-cooperative model of domestic violence in which men have all the bargaining power.

In the appendix I present a model that is a generalization of the model of Farmer and Tiefenthaler (1997) and states some of the assumptions crucial for the result that are not indicated in Farmer and Tiefenthaler (1997). The model is a straightforward Nash bargaining model in which utility is a function of consumption and violence with the man's utility is increasing in violence and the woman's decreasing in violence. The results illustrate how changes in a woman's wage affect her bargaining power and thus the level of violence by affecting her outside option. Two implications of the model are worth highlighting as they inform the empirical analysis. First, relative wages matter, not absolute wages. Second, it is the potential wage that determines one's outside option, not necessarily the actual wage (Pollak, 2005).[2] This suggests that one should focus on relative labor market conditions for women, not necessarily actual wages in this analysis. This also implies that improving labor market conditions for women will decrease violence even in households where women do not work.

In the next section I review the results of previous empirical work that examines the relationship between income and violence.

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[2] This is due to the fact that a woman's earnings at her threat point determine her bargaining power and earnings at the bargaining equilibrium do not necessarily equal earnings at the threat point. Pollak (2005) provides an example of a married woman who does not work (zero wages) at the cooperative equilibrium but who would work in the event of the dissolution of the marriage.

C. Previous Empirical Work on the Relationship Between Wages and Violence

Existing empirical research based on survey data has generally found that women with lower wages experience more violence. However, this work is limited in three respects. First, these studies fail to establish a causal relationship between domestic violence and women's wages by, for example, failing to account for omitted variable bias or reverse causality. Second, these studies focus largely on the woman's own wage when a household bargaining model suggests not only that the <u>relative</u> wage matters but that <u>potential</u>, not actual, wages determine bargaining power and levels of violence. Third, they are based on data from household surveys which are prone to non-random underreporting (Ellsberg, 2001).

The pioneering study of the relationship between women's income and violence is Gelles (1976) who finds that the fewer resources a woman has, the less likely she is to leave an abusive relationship. More recently, Bowlus and Seitz (2005) using structural estimation methods find that female employment has a large negative and significant effect on abuse. Interestingly, they also find that men are more responsive to policies designed to reduce the gains to repeat abuse than women are to policies reducing the cost of leaving violent marriages. Other work includes Macmillan and Gartner 91999) and Dugan, Nagin and Rosenfeld (1999).

Some studies have utilized panel data on women who were victims of domestic violence to examine the impact of changes in income over time on violence for a given woman (Tauchen, Witte and Long, 1991; Farmer and Tiefenthaler, 1997). While this approach enables one to overcome the potential for omitted variable bias (assuming it is time-invariant), it does not rule out the potential for reverse causality - that declines in abuse may increase a woman's productivity and earnings.

The only experimental evidence on the impact of women's economic status on domestic violence comes from a randomized intervention that combined microfinance with an education program among South African women. Women randomized to receive the intervention experienced a 55 percent drop in domestic violence relative to the control group (Pronyk et al, 2006).

Other related work on domestic violence more generally but not the relationship between violence and income include Stevenson and Wolfers (2003), Dee (2003), Fertig, Garfinkel and Mclanahan (2004), and Nou and Timmins (2005).

In the next section, I estimate the impact of women's resource on violence using individual survey data. Because of the limitations inherent in an analysis based on individual survey data, I complement it with an analysis of the impact of labor market conditions on violence against women using aggregate administrative data, overcoming some of the shortcomings of individual survey data.

III. Analysis of the Impact of Income on Domestic Violence: Individual Survey Data

A. Data and Empirical Methods

Using individual level survey data from the California Women's Health Survey (CWHS) 1998-2003, I estimate the impact of women's personal income on the probability of reporting any domestic violence. The CWHS is an annual cross sectional survey of 4000 California women. The CWHS includes a question on whether she has experienced intimate partner violence in the past year (slapping, hitting, getting beaten up, use of a weapon, threatened with a weapon).[3] The survey does not include

[3] In 2002, only a single question about experiencing any violence in the past year was included, separate questions about the type of violence were not included. As a result, the estimate of the proportion of women reporting any violence in 2002 is roughly half of what it is in 2001 and 2003. All analyses include

information about the resources of her male partner so that one cannot estimate the impact of women's relative income (as predicted by theory) but only the absolute level of her income. In addition, women's personal income in the CWHS is not provided as a continuous measure but rather as ten ranges from less than $10,000 a year to more than $100,000 per year, in increments of $10,000. To create a continuous measure each woman was assigned the midpoint of the range of income she reported, introducing substantial measurement error and attenuation bias which instrumental variable methods can reduce.

Characteristics of women in the CWHS reporting abuse and trends in abuse over this period are consistent with other survey data (see section IIA). The share of women reporting domestic violence declined from 6.2 percent in 1998 to 4.7 percent by 2003, a decline of 25 percent over five years (Appendix Table 1). In the second panel of the table is the probability of violence and mean annual income by race, education, age and marital status. Violence is highest among Black women (7.2 percent) and lowest among Asian women (3.5 percent). Violence varies inversely with education: 7.2 percent of high school drop outs report violence as opposed to 2.9 percent of college graduates. Violence also declines with age. The correlation between race, education, age and violence underscore the importance of controlling for omitted variables that may bias estimates of the impact of income on violence.

To estimate the impact of income on domestic violence in individual data, the following is estimated with a probit model:

$$DV = \delta_1 INCOME + \delta_2 RACE + \delta_3 AGE + \delta_4 EDUCATION + \gamma YEAR + \theta COUNTY + \varepsilon \qquad (1)$$

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year fixed effects and because of this change in question wording were run a second time excluding 2002, with no change in the results.

DV is an indicator equal to one if she reports that she was the victim of domestic violence in the past year. INCOME is her personal income, RACE is a vector of five race dummies (Black, White, Hispanic, Asian and Other), AGE is a vector of four age dummies (<25, 25-30, 31-39, 40-49, 50-64), EDUCATION is a vector of education dummies (HS drop out, HS graduate, College), YEAR is a vector of year dummies to control for secular trends in violence and COUNTY is a vector of six dummy variables for the six largest counties in California (Los Angeles, San Diego, Alameda, Sacramento, and Fresno) that represent 45 percent of the sample.[4] The distribution function of the error term is a standard normal.

I instrument for women's income with a measure of local labor market conditions to both overcome the bias generated by the endogeneity of women's income and to account for the fact that potential wages (as reflected in local labor market conditions) and not actual wages should affect the level of violence. To create this measure I take advantage of the history of sex-segregation by industry to construct a measure of local labor market conditions faced by women that is based on wage changes in industries dominated by women.[5] Gender and racial segregation by industry is well-established (Bayard et al, 1999; Tomasovik-Devey, 1993). For example, data for California reveal that 72% of service industry employees are women while 90% of those employed in the construction industry are men.

Average annual wages are thus calculated separately by gender and race in each county as follows:

$$\overline{w}_{grcy} = \sum_j \gamma_{grcj} w_{cyj} \qquad (2)$$

[4] Including dummy variables for each of the 58 counties in California is infeasible given the small sample sizes in most counties outside the 6 largest ones.
[5] This assumes an upward sloping labor supply curve. Fixed costs of migration would give rise to such a curve. The existence of persistent wage differences across local markets supports this assumption.

where c indexes county, g gender, r race, y year and j industry. W_{cyj} is the annual wage in industry j in county c year y from the Bureau of Economic Analysis annual survey of employers. γ_{grjc} is the proportion of women (or men) with no more than a high school degree of a given race working in industry j in county c (from the 1990 census). I focus on the low-skilled as violence is much more prevalent among this group. This proportion is fixed over this period so that changes in the wage do not reflect selective sorting across industries over this period. However, there appears to be very little redistribution between 1990 and 2000 (see Appendix Figure 1). This measure is arguably exogenous as it is driven primarily by changes in the demand for labor in industries dominated by women relative to men – not by changes in their underlying productivities which might independently affect violence. Similar measures of labor demand were used by Hoynes (2000) in her study of the impact of demand conditions on welfare participation among low income women in California.[6]

B. Results

In Table 1 are probit estimates of equation (1) and marginal effects. A woman's personal income has a negative and significant effect on the probability of experiencing any domestic violence, but the impact is small. An increase in personal income of $10,000 annually, for example, reduces domestic violence by only 2 percent. In contrast, the marginal effects of education and age are considerably larger.

However, as noted previously, it is potential wages, not actual wages, that determine one's bargaining power and the level of violence. For this reason and for issues related to endogeneity and measurement error in the income variable, I instrument for

[6] Hoynes (2000) uses county-level wages and employment in the retail and service industries as measures of the demand for low-skilled women workers.

women's income using the measure of the local demand for women's labor defined in equation (2). Estimates from the first stage are presented in Table 2. The estimated coefficient of 52 on female labor market conditions suggests that a 10 percent increase in wages paid in industries dominated by women leads to a four percent increase in female personal income. The instrumental variable probit results are presented in columns (3) and (4) of Table 1. When instrumented, the impact of a woman's own income on domestic violence increases considerably, as expected – a $10,000 annual increase in her own personal income leads to a 21 percent reduction in the probability of violence against her (a reduction of 1.1 percentage points on a baseline percentage of 5.3).

But as previously noted, results based on individual survey data are limited. To overcome these limitations I conduct an analysis based on a new source of administrative data on violence described in the next section.

IV. Analysis of the Impact of Income on Domestic Violence: Administrative Data on Hospitalizations

A. Data

For the aggregate analysis, I develop an alternative measure of violence: the number of women admitted to the hospital for an assault. This measure is derived from hospital discharge data from the state of California for 1990-2003. These data include external cause of injury data for all hospital admissions and medical personnel classify injuries as assaults, accidents or self-inflicted injuries. For each county, I calculate annual race-specific rates of hospitalization for assault for women age 15-44 for the period 1990-

2003.[7] The final dataset consists of a panel of 2261 observations (41 counties x 14 years x 4 racial groups).[8] Summary statistics are presented in Appendix Table 2.

This measure of violence against women has two drawbacks. First, it will include assaults against women not inflicted by an intimate partner. To the extent that most violence against women is perpetrated by intimates (estimates range from 76 to 87 percent) and I can control for levels of non-intimate violence in the analysis, I limit any bias from such misclassification.[9] A second drawback is that this measure will only include those assaults so severe as to require hospitalization (seven percent of injured women, according to the NVAW survey). Thus results based on hospitalizations may not be generalizeable to less severe acts of violence. In addition, because black women are both more likely to be abused and, conditional on abuse, subject to more severe violence they will be over-represented in the hospitalization data.

B. Hospitalization for Assaults 1990 -2003

The downward trend in domestic violence evident in the CWHS and the Bureau of Justice Statistics' NCV survey is likewise evident in California's hospitalization data. The rate of female hospitalization for assault declines by nearly 70 percent over this period from 39 per 100,000 to 12 per 100,000 (Figure 1A). This trend mirrors trends in intimate partner homicide in California derived from death certificates (Figure 1B). In

[7] Rates are calculated by combining these data with data on annual population counts by county, gender, race and age provided by the California Department of Finance. I limit the analysis to women age 15-44 because evidence based on surveys suggests that domestic violence is most prevalent among young women and the assault data reflect this as well. California is the largest state in the US with a population of 34 million. Of the state's 58 counties, 25 have populations in excess of 250,000 and eight have populations in excess of 1 million.

[8] Only the largest 41 counties are identifiable in the census data, and in those counties, some races are not well represented (there 35 cells with no observations.)

[9] Estimates from the NVAWS suggest 76 percent while evidence from a medical chart review of pregnant women admitted to the hospital for assault and presented by Goodwin and Breen (1990), suggests 87 percent.

addition, many of the risk factors for domestic violence identified in survey data are also apparent in the hospital discharge data. Black women, young women and poor women are all at higher risk for domestic violence and are much more likely to be admitted to the hospital for an assault than others (see Appendix Table 2). The rate at which women are admitted to the hospital for an assault is highly correlated with other measures of domestic violence such as arrests for domestic violence (0.77) and intimate partner homicide (0.89).

However, two other factors unrelated to declines in domestic violence could be responsible for the downward trend in female hospitalization for assault: declines in hospitalization and declines in violent crime. To explore the former, I display trends in hospitalization for assaults and non-assault injuries in Figures 1C and 1D. Both are declining considerably over this period, but hospitalizations for assaults proportionately more so. And to address the possibility that the decline in violent crime over this period may be responsible for the trend in female hospitalization for assault, I compare hospitalization for assault for males and females over this period in Figure 1E, assuming the decline for males captures declining rates of violent crime (Figure 1F). Over this period, assaults for males do decline, but at a slower rate (20 percent) compared to females (30 percent). While this suggests that underlying trends in both hospital utilization and violent crime explain some of the decline in the measure of domestic violence, they do not explain all. However, this underscores the need to control for such underlying trends over this period to identify the effects of the wage ratio on domestic violence, a point to which I return in the analysis.

C. Empirical Estimation Strategy

The panel structure of the administrative data enables identification of the impact of relative wages on domestic violence from changes within each county over time which implicitly controls for all differences (observed and unobserved) between counties and avoids identification from comparisons across different counties. The following equation is estimated:

$$DV_{cry} = \alpha + \beta_1 WAGERATIO_{cry} + \beta_2 UNEMP_{cy} + \beta_3 INC_{cy} + \beta_4 RACE_r + \beta_5 VIOLENCE_{cry} + \beta_6 LN(IMMIGRATION)_{cy} + \beta_7 LN(INCARCERATION)_{cry} + \gamma YEAR_y + \theta COUNTY_c + \varepsilon_{cry} \quad (3)$$

In this equation, c indexes county, r race and y year. DV refers to the measures of domestic violence derived from the hospitalization data and defined by year, county and race (for the base specification this is the hospitalization rate for assaults per 100,000 women age 15-44). WAGERATIO is the ratio of female to male earnings within race with wages constructed according to equation (2), though alternative measures (the linear difference between male and female wages and the log of the wage ratio) are also considered.[10] Using wage constructed according to equation (2), the female/male wage ratio increases six percentage points between 1990 and 2003 from 0.84 to 0.90. This measure of the wage ratio understates the true difference in earnings as it only captures wage differences due to industry segregation, failing to capture other differences due to occupational segregation, discrimination, differences in labor force participation, or differences in skill. The trend in the earnings ratio from the annual March CPS for low

[10] Examining the impact of relative wages within racial groups is justified given that inter-racial relationships are still relatively rare over this period: 14% for 18-19 year olds, 12% for 20-21 and 7% for 34-35 year olds (Joyner and Kao, 2005).

skilled (a high school degree or less) in California follows a similar upward trend over this period rising from 0.52 to 0.62.

Additional covariates that capture changes in the business cycle and trends in violent crime more generally are also included. UNEMP is the annual unemployment rate in the county and INC is per capita income in the county and year. These are included so that the impact of relative income can be identified separately from the impact of general economic conditions in the county. RACE is a vector of race dummies (Black, Asian and Hispanic – white is excluded) and their inclusion controls for the substantial differences in rates of violence across races. VIOLENCE is the non-intimate homicide rate by county, race and year and is included to control for trends in underlying violence.

To control for changes in the supply of low-skilled labor that may affect the wage ratio, I include measures of the number of new immigrants and the incarceration rate: IMMIGRATION is the number of immigrants in the county calculated by the California Department of Finance and INCARCERATION is the flow of newly incarcerated men (those incarcerated – those released). [11] Finally, year and county fixed effects are included to control for any unobserved fixed differences between counties and state-wide secular trends in domestic violence, respectively. The latter will control for all <u>state-wide</u> policy changes such as welfare reform, expansions in the EITC, changes in Medicaid eligibility or state laws regarding the prosecution of domestic violence that may affect rates of domestic violence. All regressions are weighted by cell size (female population 15-44 by race, county and year).

I argue that the weighted average female/male earnings ratio is a good measure of the local labor market conditions based by women relative to men as it primarily reflects

[11] I take the natural log of these two variables due to the difficulty defining appropriate denominators to calculate a ratio.

changes in the demand for female and male unskilled labor and not the underlying productivities of females and males in a particular labor market that may independently affect rates of violence. However, to control for other possible supply-side factors that could influence this ratio (in addition to the influx of immigrants and incarceration rates), I reconstruct the wage ratio to further limit the possibility that the ratio reflects changes in underlying productivities of women over this period. The alternative measure of wage is average wage for women (or men) in a given county*year*race cell constructed as above except that the industry wage is based on the average industry wage in the rest of the state (all counties except the given county).

$$\overline{w}_{grcy} = \sum_j \gamma_{grcj} w_{-cyj} \tag{4}$$

Constructed in this way, the measure does not reflect changes in industry wages that might be caused by changes in the county's labor supply. Identifying variation comes solely from the industrial composition of each county: counties with many workers in industries characterized by large (state-wide) wage growth will experience larger increases in average wages than counties with workers in low wage-growth industries. Identification does not arise from variation in industry-level wage growth across counties. This measure is similar to other exogenous measures of demand for labor developed by Bartik (1991) and used by Blanchard and Katz (1992) and Autor and Duggan (2003).

D. Results

Estimates of equation (3) are presented in Table 3. In the first panel of the table are estimates of the impact of the female/male wage ratio on multiple measures of assault. The coefficient estimate in the first panel of -28.43 suggests that the increase in the wage ratio from 0.84 to 0.90 over this period led to a decline in the rate of female hospitalizations of 4.4 percent. Female hospitalizations for assaults declined by 70 percent over this period, suggesting that the increase in the wage ratio accounts for 6.3 percent of this decline.

However, as previously noted the number of women admitted to the hospital for an assault also captures changes in patterns of hospitalization over this period (see Figure 1C). To account for this, I present estimates of the impact of the wage ratio on violence as measured by the share of all hospitalizations for injuries that are the result of an assault in column (2). This measure also mitigates against potential measurement error introduced by imprecision of the population counts. Again, as the wage ratio increases, the proportion of hospitalizations due to an assault declines. The observed increase in the wage ratio explains roughly 6.6 percent of the decline.

To address the concern that the decline in hospitalizations for assaults also reflects dramatic declines in violent crime more generally over this period (Figures 1E and 1F), I further refine the measure of violence. I assume that the decline in violent crime more generally over this period is reflected in the decline in hospitalizations for assaults among males. In column (3) of Table 4 I include the rate of male hospitalizations for assaults as a regressor. In column (4) I redefine the measure of domestic violence to be the share of all injuries that results from an assault for women relative to the same measure for males. This measures declines from 0.31 to 0.21 over this period, or 30 percent, very similar to the 27 percent decline in intimate partner homicides among women in California witnessed over this period. The regression coefficient of -0.173 in column (4) suggests that the decline in the wage gap of six percentage points explains ten percent of the decline in domestic violence witnessed over this period. In column (5) are estimates of the impact of the wage ratio on the ratio of female assaults to male assaults, controlling for the ratio of female non-assault injuries to male non-assault injuries.   The estimate is very similar to that in column (4).

E. Robustness

Redefining the Wage Gap

To explore whether the results are sensitive to the definition of the wage gap, I redefine the wage gap to be the log of the wage ratio and the linear difference of male and female wages in the second and third panels of Table 4, respectively. Regardless of the way that the wage difference is measured, the results are qualitatively the same: closing the wage gap leads to a decline in the rate of female hospitalization for assaults.[12]

Violence against Men

To compare how the wage ratio affects female assaults compared to male assaults, I redefine the outcome measure to be the natural log of assaults. In the first column of Table 4A, I present estimates of the impact of the wage gap on the natural log of female hospitalizations for assault, and in the second column I do the same for the natural log of male assaults. For these regressions, because of small cell with zero hospitalizations for assault, I restrict the analysis to cells with at least 15,000 women (or men). These results suggest that a six percentage point increase in the female/male wage gap would lead to a 7.8 percent decline in female hospitalizations for assault but only a 2.2 percent decline in male hospitalizations for assault.[13] In the third column, I present the results of a regression of the impact of the wage gap on female hospitalizations controlling for male

[12] When male and female wages are entered separately, women's wages reduce the rate of female assaults (coefficient -.161 that is statistically significantly different from 0 at the 1 percent level) while male assaults have a positive but insignificant effect on assaults (coefficient of .106).
[13] The 2.2 percent decline among men may reflect reductions in the number of women in battering relationships who assault their male partners in self-defense as suggested by some criminologists.

hospitalizations. The results suggest that an increase in the wage gap of 6 percentage points leads to a 6.7 percent decline in violence.

Zipcode-Level Analysis

One possible reason why we find that as increase in women's wages is associated with declines in violence is that as women earn more, they move to "safer" neighborhoods and thus are less likely to be the victim of crime. To rule this possibility out, I create zipcode level measures of violence and examine the impact of changes in the wage ratio on violence within zipcode. Because population by race, gender and age is not available at the zipcode level (except from the decennial census), the outcome examined is the ratio of female assaults to male assaults (as in column 5 of Table 4).

The results (Table 4B) are not directly comparable to the previous analysis because the five digit zipcode is only available for years 1991 and 1994-2000 in the hospital discharge data and is missing for 25,000 discharges. This non-random sampling results in important differences in the outcome measure: for the zipcode sample, the ratio of female to male assault is .076 whereas in the full sample the ratio is .131 and the decline in violence is also smaller, from .094 to .061. Thus, even though the coefficient estimates are one third the value of estimates based on the full sample, the interpretation is similar: the increase in the ratio of female to male wages still explains ten percent of the decline in violence over this period.

Falsification Tests

I also estimate the impact of the wage difference on two outcomes for which I expect small or no effects: female hospitalizations for attempted suicide and car crashes.

There does not appear to be any significant effect of the wage ratio on either of these measures.[14]

Exogeneity of Wages

To address the possibility that the measures of the female- male wage gap may not just reflect an increase in demand for female labor but an increase in the productivity of local area women which may be correlated with propensity for domestic violence, I use county level wage gap as calculated in equation (4). Recall that this measure is based on the industrial structure of the county, but the wage in the rest of the state (excluding the county). These estimates are presented in Table 5. The estimates are precise and similar to the estimates based on the previous county-level measure of wages.

Exposure Reduction

The findings thus far provide evidence in favor of a marital bargaining model in which an increase in women's relative wage increases her bargaining power, thereby decreasing violence against her. However, these findings do not rule out the possibility of an alternative explanation - exposure reduction. Because an increase in women's wages is likely to be accompanied by an increase in female employment, finding that violence falls as wages rise may be evidence of either a bargaining story or exposure reduction. In order to test whether exposure reduction is responsible for these findings, I estimate the impact of changes in the wage ratio on assaults that occurred during the weekday vs. the weekend. If exposure reduction explains the findings then I should see a larger decrease in assaults during the weekday than the weekend. Only 5 years of data (1990, 1992, 1993, 1995 and 1996) include information on day of week of admission and are used for this analysis.

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[14] The coefficient estimate for suicide is 0.0083 with a standard error of 0.0152; the coefficient estimate on car crashes is -0.0067 with a standard error of 0.0199.

In Table 6 I present estimates based on these five years of data. In the first two columns I present estimates of the impact of the female/male wage ratio on the share of hospitalizations for an injury that are the result of an assault on weekends (column 1) and weekdays (column 2). In the next two columns I present estimates of the impact of the linear difference in wage rates. Most of the decline in violence resulting from an increase in the wage ratio occurs during the weekend, which I argue is inconsistent with the exposure reduction hypothesis.

Based on the finding that an increase in women's relative wages reduces violence against her, I revisit previous work establishing a positive relationship between women's share of household resources and child health and well-being. While previous work has largely assumed that this relationship is attributable to increases in women's material investment in children, in the next section I explore whether an alternative mechanism (reductions in violence) may also play a role.

V. Women's Wages, Violence and Child Health

A. Background

A marital bargaining model that incorporates children yields important predictions regarding women's income and the allocation of household resources to children. In bargaining models, if mothers exert stronger preferences for their children than do fathers, then as women's income (and bargaining power) increase, household allocations to children should likewise increase. Previous empirical work on intra-household allocation has largely supported bargaining models over common preference models. Seminal work by Thomas (1990) based on survey data from Brazil found that unearned income in the hands of a mother has a bigger effect on her family's health than income under the control of a father. The positive relationship is often attributed to increased expenditures on children: given equal increases in maternal and paternal income, the former results in larger expenditures on children than the latter. More recent work by Duflo (2000) found that an increase in pension payments among women in South African households led to improvements in the health of girls in the household (as measured by height and weight for height), but not boys. In contrast, an increase in pension payments among men did not have any affect on the health of children.

The mechanism behind this relationship, however, is not well-established. Previous work linking the distribution of resources in the household to improvements in child health often assumes that an increase in material investment is responsible. Some evidence that a reallocation of resources from the father to the mother results in an increase in material investments in children is provided by Lundberg, Pollak and Wales (1997). They find, based on data from the UK, that an exogenous increase in maternal income leads to an increase in expenditures on women and children's clothing.

The relationship between women's relative resources and violence established in the first part of this paper, however, suggests that reductions in violence provide an additional explanation for why an increase in women's relative income results in improved child outcomes. This mechanism has not previously been considered. In the rest of this paper I explore the relationship between violence and birth outcomes and provide the first estimates of a causal relationship.

B. Previous Literature on Violence and Birth Outcomes

Previous studies have provided estimates of the prevalence of domestic violence among pregnant women in the US that range from 0.9% to 20.1% (Gazmararian, et al

1996).[15] A number of studies have found that violence often initiates or escalates during pregnancy (Stewart and Cecutti, 1993; Helton, McFarlane and Anderson, 1987; Amaro, Fried, Cabral and Zuckerman, 1990). Psychologists have offered one possible explanation for the increase in violence during pregnancy: sexual jealousy inspired by the uncertainty of paternity. In an interview of 258 men convicted of spouse abuse, Burch and Gallup (2004) found that the frequency and severity of abuse directed toward pregnant partners was double that directed toward partners who were not pregnant and that sexual jealousy was also greater for men with pregnant partners.

Medical studies have documented a negative correlation between domestic abuse during pregnancy and birth outcomes. Valladeras (2002) found that 22 percent of mothers with low birth weight (LBW) infants experienced physical abuse as opposed to five percent of full weight infants, controlling for potential confounders such as age, parity, socio-economic status and smoking. In a meta-analysis of eight studies, Murphy el al (2001) found that women who reported abuse during pregnancy were more likely than nonabused women to give birth to a baby with LBW (OR 1.4).

Violence affects pregnancy outcomes via multiple mechanisms. Abuse resulting in blunt trauma to the maternal abdomen can cause abruptio placentae, fetal fractures, rupture of the maternal uterus, liver, spleen and antepartum hemorrhage. Non-abdominal trauma can also cause uterine contractions, premature rupture of membranes and infection. Finally, abuse may also lead to the exacerbation of chronic illnesses such as hypertension, diabetes or asthma which can negatively affect the fetus.

[15] Examples include Hillard (1985) who found that 3.9 % of 742 prenatal women reported abuse during pregnancy; Helton, McFarland and Anderson (1987) found that 8% of 290 pregnant women reported violence. Berenson, Stiglich, Wilkinson and Anderson (1991) found that 5.5% reported abuse during the current pregnancy and in a postpartum sample of 488 women, Campbell, Poland, Waller and Ager (1992) found that 8.3% of the women reported abuse during the pregnancy.

While biological pathways linking violence to poor birth outcomes exist, previous studies have not effectively isolated the impact of violence on birth outcomes from other maternal characteristics such as poverty and risk taking behavior (smoking, drinking) that are correlated with both violence and birth outcomes. I use propensity score and non-linear instrumental variable methods to establish the first causal estimate of the impact of violence on low birth weight.

C. Data on Violence during Pregnancy and Birth Outcomes

To estimate the impact of violence on birth outcomes I use a unique dataset that links maternal hospitalizations in the nine months prior to birth with detailed natality data that includes information on birth outcomes from California for the period 1991-2002 (excluding 1998).

Of the more than 5 million births over this period, only 1656 women were admitted to the hospital for an assault while pregnant – roughly 3 per 100,000. However, the rate is much higher among disadvantaged women: 50 per 100,000 for those on Medicaid and 164 per 100,000 black women. Sample means for these data presented in Table 7 columns (1) and (2) illustrate how women who are admitted to the hospital for an assault are more likely to suffer worse birth outcomes and are more likely to come from disadvantaged backgrounds (poorer, less educated, younger and more likely to be black) and engage in risky behavior such as using drugs and smoking which may independently affect birth outcomes.[16] In columns 3 and 4 of Table 7 are average birth outcomes and maternal/paternal characteristics of women who suffered unintentional injuries and car crashes. These women suffer worse birth outcomes than women with no injuries, but not

[16] Smoking and drinking variables are under-reported and measured with considerable error in the California natality data because unlike most states, California only requires reporting if the behavior resulted in a complication.

as bad as those who have been assaulted. In addition, they are not nearly as disadvantaged as victims of assault, suggesting significant negative selection into violent relationships.

While these data represent the universe of California births, they exclude women who miscarried or aborted. How this might bias estimated effects depends on from what part of the distribution we believe these women are drawn. One might reasonably argue that these women suffer (or expect to suffer) the most extreme violence and the worst birth outcomes, suggesting that estimates that exclude these women will be biased downward.

D. Empirical Estimation Strategy and Results

To estimate the impact of violence on birth outcomes, I first estimate probit models of the impact of admission to the hospital for an assault on the probability a child is born LBW including controls for maternal background. This is followed by propensity score matching estimates and bivariate probit estimates that account for the non-random selection into violent relationships.

The probit model of the impact of violence on the probability a child is born LBW is as follows: let the indicator $LBW_i=1$ if mother I gives birth to a LBW infant and $LBW_i=0$ otherwise. The birth production function is described by the latent variable model:

$$LBW_i^* = \beta X_i + \delta V_i + \varepsilon_i$$

Where LBW_i^* is the underlying health of the child, X_i is a vector of individual maternal characteristics and V_i is an indicator for whether the birth mother was admitted to the hospital for assault while pregnant. The probability that a baby is born LBW is Prob $[LBW_i=1]= \Phi [\beta X_i + \delta V_i]$ where Φ is a standard normal cdf.

In Table 8 column (1) and (2) are probit estimates and marginal effects of the impact of assault on LBW.[17] Women who are assaulted are 3.7 percent more likely to have a LBW infant. This effect is considerably larger than that of other maternal characteristics that have been shown to affect birth outcomes (such as poverty and maternal education) with the exception of being black which has roughly the same negative impact on birth weight as assault.

For the propensity score/matching estimator, I match women who were assaulted with women who were not based on propensity score methods.[18] The estimates derived from propensity score matching methods suggest that women who are the victim of violence are 8.5 percent more likely to have a LBW birth relative to similar women who have not been assaulted and their babies weigh on average 180 grams less. The estimates are significant at the 1 percent level. That this estimated effect is larger than the effect based on the probit model is attributable to the weighting scheme employed in matching: women most likely to be assaulted (ie, have the highest propensity score) receive a higher weight in the matching estimate. Figure 2 which displays matching estimates of the impact of assault on LBW by propensity score illustrates this point. The estimated impact of assault is much higher for women with higher propensity scores – as high as 22 percent. If estimates based on these women receive greater weight in the matching estimator, then this would explain the source of difference in the probit and matching estimates. One can argue that an estimate that reflects the impact of violence on birth

[17] Estimates of the impact of violence on infant mortality were positive but not significant.

[18] I estimated the probability of assault (propensity score) based on all control variables included in Table 12, then matched women who were assaulted with those who were not with a very similar propensity score (within the same bin – for 150 bins). Matching was done without replacement and the matching estimator was weighted by the propensity score.

outcomes for those most likely to be the victim of violence is more useful than one that gives equal weight to women at very low risk of abuse.

As a final estimation strategy, I estimate the impact of assault on the probability of LBW using a bivariate probit model. In this model, the birth production function is described by the latent variable model as before, but now suppose that the process by which women are hospitalized for an assault is described by the latent variable model:

$$V_i^* = \tau X_i + \gamma Z_{cry-1} + \mu_i$$

where V_i^* is the amount of violence, X_i is a vector of individual characteristics and Z_{cry-1} is a variable not contained in X_i that affects violence against women. To allow for the possibility that the unobserved determinants of birthweight are correlated with unobserved determinants of violence, assume that $E[\varepsilon_i] = E[\mu_i] = 0$, var $[\varepsilon_i]$ = var $[\mu_i] = 1$ and $\text{cov}[\varepsilon_i, \mu_i] = \rho$. The instrument for violence in this case (Z_{cry-1}) is a measure of the strictness of prosecutorial policies towards domestic violence in the previous year and is defined at the county-year-race level. In California, as elsewhere, laws regarding the prosecution of domestic violence are determined at the state level. However, prosecution of domestic violence falls to the local (county) prosecutors offices. Local county prosecutors in California have wide discretion over the prosecution of spousal assault. They vary in terms of whether they have separate offices/prosecutors who specialize in domestic violence, the amount of training their prosecutors receive and the presence of advocates assigned to women bringing charges. Unfortunately, data on police and prosecutorial policies are not available on a consistent basis.[19] Instead, I proxy for prosecutorial policies by calculating the proportion of all men arrested for domestic violence who are sentenced to jail for each race in each county and year. If either the

[19] Data on such policies are only available for the 7 largest counties in California and only up until 1996.

deterrent or incapacitation effects of incarceration are strong (Levitt, 1996), the proportion of offenders who go to jail in the previous year (referred to here as the lagged incarceration rate) may serve as an appropriate instrument for the level of violence witnessed today: as the incarceration rate in the previous period increases, violence should decline.

Table 8 columns 3-5 contain coefficient estimates and marginal effects from the bivariate probit model. In column (3) are estimates of the impact of violence on the probability LBW, column (4) contains the marginal effects and in column (5) are coefficient estimates of the determinants of violence. Increasing the percentage of men who go the jail for domestic violence conditional on arrest for domestic violence significantly decreases the probability that a pregnant woman will be admitted to the hospital for an assault in the next year. Over this period, the incarceration rate increased from 20 percent to 74 percent. An increase of this magnitude leads to a very small (2 percent) decline in the probability of assault on average, a point to which I return.

Based on estimates from the bivariate probit, women who are admitted to the hospital for an assault face an 18 percent increase in the probability of LBW birth – a considerably larger effect than was found via probit and also larger than estimates from the matching/propensity score methods. To understand this discrepancy, I examine whether there is heterogeneity in the effect of the lagged incarceration rate (Z_{cry-1}) on violence. In figure 3 are coefficient estimates of the impact of Z_{cry-1} on whether the woman was hospitalized for assault by propensity score. It appears that the lagged incarceration rate is most negatively related to assault among those with the highest propensity score for assault. We know from Figure 2 that the impact of assault on birthweight is likewise higher among those with the highest propensity score, suggesting that a local average treatment effects (LATE) interpretation can explain the increase in magnitude of the bivariate probit estimates over the probit and propensity matching score estimates. In sum, the results presented here suggest that assault during pregnancy leads to an increase in the probability of a low birth weight birth of 4 to 18 percent. This range reflects considerable heterogeneity in the impact with the children of those women most disadvantaged and most likely to be the victim of violence suffering the most.

VI. Conclusion

Over the past fifteen years, violence against women has declined as their employment and earnings have increased. A model of household bargaining that incorporates violence is consistent with these trends. I find empirical support for a causal relationship between labor market conditions for women and violence using both individual survey data and administrative data: recent improvements in labor market conditions for women relative to men have led to a ten percent reduction in violence against them. This finding suggests that in addition to more equitable redistribution of resources, policies that serve to narrow the male-female wage gap also reduce violence and the costs associated with it. The estimates imply that if women were equally represented in high wage industries erasing this source of the wage gap, violence against them would decline by an additional 16 percent.

Based on this finding, I revisit previous work establishing a positive relationship between women's share of household resources and child health and well-being. While previous work has largely assumed that this relationship is attributable to increases in women's material investment in children, I provide evidence that a reduction in violence, at least in the case of birth weight, is also an important determinant. These results suggest that in addition to addressing concerns of equity, improved pay parity can have positive effects on the health of American women as well as important intergenerational effects.

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Appendix I: A Model of Household Bargaining with Violence

In this appendix I develop a simple model of household bargaining that incorporates violence and shows under what assumptions an increase in women's relative income leads to a decline in violence.

Let $U_w\ (C_w, S)$ be a woman's utility which is increasing in her own consumption (C_w) and increasing in safety (S) and let $U_m\ (C_m, V)$ be a man's utility which is increasing in his own consumption (C_w) and in violence (V). Assume that there is an upper bound to violence (death) $\overline{V}$ and $S = \overline{V} - V$. Both utility functions are assumed to be strictly concave, monotonically increasing, differentiable and homothetic. I is total household income and α is the share of income a woman would get if she were not in the partnership. To focus on partnerships that will experience some violence, assume that the marginal rate of substitution between consumption and violence is greater for women than men where V=0 ($MRS_m\ ((1-\alpha)I, 0) < MRS_w\ (\alpha I, \overline{V}\ ))$. We denote as T the set of feasible utility pairs (U_m, U_w) that the partners may obtain if they reach an agreement and the utility each gets from dissolution of the partnership (the disagreement payoff or single state utility) as $(d_m, d_w) = (Um(1\text{-}\alpha)I, 0), Uw\ (\alpha I, \overline{V}\ )$.

I first show that the domestic violence problem constitutes a Nash bargaining problem and that a Nash bargaining solution provides a unique solution to the problem.

Lemma 1: Under the above assumptions, the problem is a Nash bargaining problem with a Nash bargaining solution (Nash, 1950).

Proof:

The following necessary and sufficient conditions are met:

1) The set T is compact. This is satisfied since the set of feasible allocations is compact and U_m and U_w are continuous.

2) The set T is convex. This is satisfied since the set of feasible allocations is convex and U_m and U_w are continuous.

3) $d= (U_m((1- \alpha)I, 0), U_w(\alpha I, \overline{V}))$ is a member of T.

4) For some member (U_m, U_w) of T it is the case that $U_m>d_m$ and $U_w>d_w$. This holds under the assumption that $MRS_m((1- \alpha)I, 0) < MRS_w(\alpha I, \overline{V})$. ■

To allow for different degrees of bargaining power between partners, we will use Kalai's asymmetric Nash bargaining solution. The solution to the asymmetric bargaining problem (U^*_m, U^*_w) maximizes the following expression as shown by Kalai (1983):

$(U_m-d_m)^{\tau} (U_w-d_w)^{1-\tau}$

subject to $(U_m, U_w) \geq (d_m, d_w)$ and (U_m, U_w) feasible, where τ is a measure of bargaining power of the man.

I study next how an increase in the relative wage of women affects utilities of both men and women.

Lemma 2: An increase in α results in an increase in U^*_w and a decrease in U^*_m.

Proof:

Take α and α' such that $\alpha' > \alpha$. Then $d_w(\alpha') > d_w(\alpha)$ and $d_m(\alpha') < d_m(\alpha)$. Therefore,

$$\frac{\tau}{(1-\tau)} \frac{(U_w - d_w(\alpha'))}{(U_m - d_m(\alpha'))} < \frac{\tau}{(1-\tau)} \frac{(U_w - d_w(\alpha))}{(U_m - d_m(\alpha))}$$

and by convexity of T we have that $U^*_w(\alpha') \geq U^*_m(\alpha)$. ■

This alone does not guarantee that an increase in α results in a decline in violence. This depends on the shape of the contract curve.

Lemma 3: the contract curve has a positive slope.

Proof:

Let (C_m, V) and (C_m', V') be points on the contract curve such that $C_m' > C_m$ and $V' < V$. By definition, $MRS_m(C, V) = MRS_w(C, \overline{V} - V)$ since (C_m, V) is a point on the contract curve. This implies that $\frac{C'_m}{V'} > \frac{C_m}{V}$ and $\frac{C'_w}{\overline{V} - V'} < \frac{C_w}{\overline{V} - V}$

By homotheticity and strict concavity of utility functions, $MRS_m(C'_m, V') < MRS_m(C_m, V)$ and

$MRS_w(C'_m, \overline{V} - V') > MRS_w(C_m, V)$. This implies that $MRS_w(C', \overline{V} - V') > MRS_m(C', V')$ and the point ($C'm$, V') cannot be on the contract curve. ■

Theorem 1: An increase in α results in a decrease in violence.

Proof:

Take α and α' such that $\alpha' > \alpha$.

Then $U^*_w(\alpha') > U^*_w(\alpha)$, by lemma 2, which implies that $V(\alpha') < V(\alpha)$, since the contract curve has a positive slope by lemma 3. ■

The above simple model of household bargaining shows that under certain reasonable assumptions (namely strict concavity, differentiability and homotheticity of utility functions) an increase in a woman's income leads to a decline in violence against her.[20]

[20] Incorporating the marriage market should not affect this comparative static result. Rather, the marriage market simply influences the disagreement payoff (d) which will become the maximum of an individual's single state utility and the expected utility from another match. In this paper I do not focus on the marriage market because there is no data for California on marriages and divorces.

Table 1: Probit and IV Probit Estimates and Marginal Effects: Impact of Women's Income on Probability of Domestic Violence in Past year

	Probit		IV Probit	
	Coefficient	Marginal Effect	Coefficient	Marginal Effect
Annual Personal Income/10000	-0.010	-0.001	-0.108	-0.011
	[-1.93]		[-1.77]	
White	-0.300	-0.028	-0.145	-0.015
	[-3.39]		[-1.15]	
Black	-0.133	-0.011	-0.041	-0.004
	[-1.25]		[-0.30]	
Hispanic	-0.256	-0.021	-0.119	-0.012
	[-2.84]		[-0.97]	
Asian	-0.482	-0.031	-0.314	-0.025
	[-4.35]		[-2.12]	
<HS	0.124	0.012	0.043	0.004
	[2.58]		[0.65]	
HS	0.088	0.008	-0.012	-0.001
	[2.27]		[-0.18]	
Age <=25	0.788	0.115	0.639	0.093
	[12.35]		[4.74]	
Age 25-30	0.724	0.103	0.656	0.097
	[10.92]		[7.25]	
Age 31-39	0.580	0.067	0.516	0.064
	[9.26]		[6.75]	
Age 40-49	0.370	0.039	0.334	0.039
	[5.85]		[5.06]	
Age 50-64	0.018	0.002	0.011	0.001
	[1.35]		[0.78]	
1999	0.062	0.006	-0.028	-0.003
	[1.23]		[-0.64]	
2000	-0.019	-0.002	0.017	0.002
	[-0.36]		[0.30]	
2001	-0.072	-0.006	0.092	0.010
	[-1.34]		[1.11]	
2002	-0.527	-0.036	0.072	0.008
	[-8.21]		[0.50]	
2003	-0.142	-0.012	0.094	0.010
	[-2.63]		[1.23]	
Los Angeles	-0.056	-0.005	0.026	0.003
	[-1.40]		[0.26]	
San Diego	0.012	0.001	0.043	0.004
	[0.22]		[0.83]	
Alameda	0.049	0.005	0.010	0.001
	[0.62]		[0.19]	
San Francisco	-0.018	-0.002	-0.199	-0.018
	[-0.14]		[-1.99]	
Sacramento	0.100	0.010	-0.643	-0.047
	[1.29]		[-7.65]	
Fresno	0.035	0.003	-0.271	-0.024
	[0.35]		[-2.96]	
Constant	-1.800		-1.435	
	[-16.48]		[-3.92]	
Observations	18636		18636	

Z statistics in brackets below coefficient estimates

Table 2: First Stage Estimates of Impact of Female Labor Market Conditions on Women's Annual Income - CWHS

Female Labor Market Conditions	52.206
	[11.35]
White	0.233
	[1.29]
Black	-0.094
	[-0.47]
Hispanic	0.320
	[1.77]
Asian	0.695
	[3.54]
<HS	-0.651
	[-9.06]
HS	-0.773
	[-13.95]
Age <=25	-1.277
	[-16.37]
Age 25-30	-0.413
	[-5.06]
Age 31-39	-0.275
	[-3.87]
Age 40-49	-0.036
	[-0.52]
Age 50-64	-0.047
	[-2.37]
1999	0.189
	[3.45]
2000	-0.037
	[-0.47]
2001	0.211
	[1.89]
2002	0.451
	[2.55]
2003	-0.173
	[-1.55]
Los Angeles	-0.002
	[-0.01]
San Diego	-0.136
	[-1.78]
Alameda	0.163
	[2.11]
San Francisco	-1.509
	[-19.49]
Sacramento	-1.426
	[-18.73]
Fresno	-1.429
	[-18.75]
Constant	2.860
	13.12
Observations	18636

Z statistics in brackets below coefficient estimates

Table 4: Estimates of the Impact of Labor Market Conditions on Violence

	(1)	(2)	(3)	(4)	(5)
	Female assaults	**Female Assaults/ Total Injuries**	**Female Assaults/ Total Injuries**	**(Fem Assault/Injuries)/ (Male Assault/Injuries)**	**Female/ Male Assaults**
Wage Ratio					
Female/Male Wage	-28.431	-0.0339	-0.0208	-0.1729	-0.1705
	[9.1036]	[0.0093]	[0.0089]	[0.0496]	[0.0491]
Male Assaults			0.1218		
			[0.0156]		
Female/Male Non Assaults					0.0575
					[0.0286]
Black	47.5185	0.063	0.0438	0.0777	0.0739
	[4.5636]	[0.0046]	[0.0046]	[0.0179]	[0.0183]
Hispanic	-9.6552	0.0161	0.0014	-0.1067	-0.0959
	[1.6139]	[0.0015]	[0.0023]	[0.0076]	[0.0085]
Asian	-8.8909	0.0019	-0.0036	0.0096	-0.0142
	[0.6324]	[0.0015]	[0.0015]	[0.0136]	[0.0123]
Ln(per capita income)	-5.8704	0.0101	0.008	-0.0948	-0.079
	[10.0347]	[0.0102]	[0.0097]	[0.0527]	[0.0521]
non-intimate homicide rate	93,127.69	85.2876	61.6599	-66.5776	-68.0813
	[9,852.7269]	[8.9150]	[7.9142]	[22.4166]	[22.5438]
unemployment rate	-36.6822	-0.0595	-0.0562	0.1497	0.0932
	[40.6674]	[0.0494]	[0.0479]	[0.4162]	[0.4213]
Ln(immigration)	4.6829	0.0071	0.0064	-0.0096	-0.0088
	[1.9165]	[0.0022]	[0.0021]	[0.0156]	[0.0155]
Observations	2261	2261	2261	1865	1853
R-squared	0.74	0.78	0.79	0.25	0.26
Ln(wage ratio)					
ln(female wage/male wage)	-22.9341	-0.0293	-0.0179	-0.1492	-0.1481
	[7.5461]	[0.0080]	[0.0077]	[0.0428]	[0.0423]
Male Assaults			0.1217		
			[0.0155]		
Observations	2229	2229	2229	1861	1851
R-squared	0.74	0.78	0.79	0.25	0.26
Linear difference of wages					
Male wage-female wage	0.0407	0.0001	0.0001	0.0007	0.0007
	[0.0351]	[0.0000]	[0.0000]	[0.0002]	[0.0002]
Male Assaults			0.1232		
			[0.0156]		
Observations	2261	2261	2261	1865	1853
R-squared	0.77	0.78	0.79	0.25	0.26

Each observation is a county-year-race cell.
All regressions include county and year dummy variables. All regressions weighted by population.
Robust standard errors clustered on county.

Table 4A: Estimates of the Impact of Labor Market Conditions on Assaults - Natural Logs

	Ln(female Assaults)	Ln(Male Assaults)	Ln(Female Assaults)
female wage/male wage	-1.2965	-0.3707	-1.1119
	[0.3245]	[0.1923]	[0.3067]
Male Assaults			0.3353
			[0.0635]
Black	1.5026	1.3507	1.1918
	[0.1119]	[0.0806]	[0.1304]
Hispanic	0.4359	1.2056	0.0231
	[0.0736]	[0.0570]	[0.1021]
Asian	0.094	0.8319	0.0082
	[0.1271]	[0.0966]	[0.1196]
Ln(per capita income)	0.0344	0.0165	-0.049
	[0.2848]	[0.1637]	[0.2776]
non-intimate homicide rate	545.6678	617.4327	215.9786
	[105.5956]	[86.4992]	[97.6656]
unemployment rate	0.4021	-0.7642	0.9405
	[1.4372]	[0.9887]	[1.4140]
Ln(immigration)	-0.0382	-0.0175	-0.0359
	[0.0647]	[0.0433]	[0.0625]
nonratio	-0.2855	0.1906	-0.1689
	[0.1691]	[0.1283]	[0.1383]
ln(population)	0.14	-0.2659	0.1842
	[0.0842]	[0.0648]	[0.0777]
ln(non-assault injuries)	0.8921	1.3436	0.5259
	[0.0806]	[0.0701]	[0.1049]
Observations	965	1060	963
R-squared	0.96	0.98	0.96

Robust standard errors in brackets

Each observation is a county-year-race cell.
All regressions include county and year dummy variables. All regressions weighted by population.
Robust standard errors clustered on county.

Table 4B: Impact of Relative Wages on Ratio of Female to Male Assaults - zipcode level

female wage/male wage	-0.056		
	[0.027]		
ln(female wage/male wage)		-0.047	
		[0.024]	
Male wage-female wage			0.00023
			[0.00014]
Female/Male Non Assault Injuries	0.008	0.008	0.008
	[0.002]	[0.002]	[0.002]
Black	-0.06	-0.061	-0.06
	[0.010]	[0.009]	[0.010]
Hispanic	-0.077	-0.077	-0.076
	[0.006]	[0.006]	[0.006]
Asian	-0.118	-0.118	-0.115
	[0.007]	[0.007]	[0.008]
Ln(per capita income)	-0.019	-0.02	-0.024
	[0.042]	[0.042]	[0.041]
non-intimate homicide rate	78.105	78.494	79.581
	[16.720]	[16.713]	[16.703]
unemployment rate	0.159	0.16	0.159
	[0.345]	[0.345]	[0.345]
Ln(immigration)	0.002	0.002	0.001
	[0.012]	[0.012]	[0.012]
Observations	55074	55064	55074
R-squared	0.08	0.08	0.08

Robust standard errors in brackets

Each observation is a zipcode-race-year cell. Regressions weighted by number of total hospitalizations in each cell.

Note: zipcode fixed effects included; only years 1991 and 1994-2000 included in this analysis due to lack of 5 digit zipcode in years 1990, 1992-1993 and 2001-2003

Table 5: Estimates of the Impact of Labor Market Conditions on Violence - Alternative Measure of Wage

	(1)	(2)	(3)	(4)	(5)	(6)
	Female assaults	**Female Assaults/ Total Injuries**	**Female Assaults/ Total Injuries**	**(Fem Assault/Injuries)/ (Male Assault/Injuries)**	**Ln(Female Assaults)**	**Ln(Female Assaults)**
Female/Male Wage (State wages)	-36.6386	-0.0398	-0.023	-0.2444	-1.4714	-1.1246
	[14.6451]	[0.0125]	[0.0121]	[0.0605]	[0.3820]	[0.3804]
Black	47.8359	0.064	0.044	0.0822	1.288	0.4275
	[4.6510]	[0.0047]	[0.0047]	[0.0185]	[0.1297]	[0.0738]
Hispanic	-10.4105	0.0141	-0.0009	-0.1035	0.0916	-0.2234
	[2.0428]	[0.0018]	[0.0025]	[0.0092]	[0.0959]	[0.0708]
Asian	-9.376	0.0053	-0.0005	-0.014	0.2287	-0.2418
	[0.7604]	[0.0018]	[0.0019]	[0.0126]	[0.1139]	[0.0663]
Ln(per capita income)	-7.2764	0.0062	0.0049	-0.1027	-0.0909	0.1155
	[10.2762]	[0.0101]	[0.0096]	[0.0545]	[0.2721]	[0.2594]
non-intimate homicide rate	92,845	84.9313	60.3631	-75.9704	166.7073	92.7411
	[9,882]	[9.0958]	[8.0349]	[22.7912]	[95.8800]	[91.2645]
unemployment rate	-43.7775	-0.0616	-0.0528	0.0451	0.6365	1.8124
	[39.6944]	[0.0478]	[0.0465]	[0.4458]	[1.3561]	[1.5451]
Ln(immigration)	4.9527	0.0075	0.0066	-0.0075	-0.0274	-0.0266
	[1.9767]	[0.0023]	[0.0021]	[0.0164]	[0.0619]	[0.0640]
Non assault injuries female/Non Assault inju	1.5232	-0.007	-0.0072	0.0611	0.0993	-0.696
	[1.1165]	[0.0025]	[0.0030]	[0.0302]	[0.2207]	[0.2652]
Male Assaults/total injuries			0.1266			
			[0.0158]			
ln(non-assault injuries female)					0.379	
					[0.2059]	
Ln(Male Assaults)					0.2909	
					[0.0616]	
ln (non assault injuries female)/ ln (non assault injuries male)						1.5895
						[0.2101]
Observations	2118	2118	2118	1853	963	963
R-squared	0.74	0.79	0.8	0.26	0.96	0.6

Robust standard errors in brackets

Table 6: Estimates of Impact of Labor Market Conditons on Female Assaults/Female Injuries
Weekend vs. Weekday Admissions

	Female Wage/Male Wage		Male Wage-Female Wage	
	Weekend	Weekday	Weekend	Weekday
Relative Wage	-0.137	-0.056	0.0006	0.0002
	[0.060]	[0.018]	[0.000]	[0.000]
Black	0.15	0.132	0.153	0.132
	[0.014]	[0.006]	[0.016]	[0.006]
Hispanic	0.035	0.026	0.038	0.027
	[0.009]	[0.004]	[0.010]	[0.005]
asian	0.095	0.05	0.105	0.054
	[0.029]	[0.009]	[0.028]	[0.009]
Ln(per capita income)	-0.093	-0.116	-0.059	-0.007
	[0.048]	[0.021]	[0.130]	[0.036]
ln(non-intimate homicide)	0.014	0.004	0.013	0.003
	[0.005]	[0.002]	[0.004]	[0.002]
unemployment rate	-0.36	0.056	-2.695	0.042
	[0.208]	[0.054]	[0.732]	[0.151]
Ln(immigration)	0.009	0.01	-0.007	0.007
	[0.010]	[0.004]	[0.012]	[0.004]
Ln(incarceration)	0.002	0.006	0.003	0.007
	[0.010]	[0.002]	[0.010]	[0.002]
Observations	718	816	718	816
R-squared	0.38	0.78	0.4	0.78

Robust standard errors in brackets

Table 7: Birth Outcomes and Maternal/Paternal Characteristics

	No assault	Assault	Unintentional Injury	Car Crash
Birth outcomes				
LBW	0.064	0.149	0.101	0.080
Fetal Death	0.006	0.010	0.005	0.004
Infant death	0.006	0.012	0.011	0.008
Pregnancy				
Drug use	0.001	0.009	0.004	0.001
Tobacco	0.020	0.081	0.046	0.037
Bleeding	0.005	0.009	0.006	0.010
Maternal Characteristics				
Teenage	0.12	0.22	0.13	0.16
Over 35 years old	0.10	0.05	0.10	0.08
<HS	0.34	0.44	0.31	0.28
HS	0.30	0.37	0.35	0.38
Some college	0.19	0.14	0.21	0.22
College	0.17	0.05	0.12	0.12
Medicaid	0.43	0.70	0.52	0.48
Black	0.08	0.43	0.19	0.17
White	0.91	0.56	0.81	0.82
Hispanic	0.53	0.34	0.39	0.43
Paternal Characteristics				
Black	0.08	0.40	0.18	0.18
Hispanic	0.54	0.45	0.44	0.45
White	0.36	0.14	0.35	0.35
<HS	0.30	0.28	0.22	0.21
HS	0.30	0.39	0.38	0.38
Some college	0.16	0.10	0.15	0.17
College grad	0.17	0.03	0.12	0.13
Observations	5397529	1656	3699	5270

Table 8 Impact of Prenatal Assault on LBW: Probit and Bivariate Probit Estimates

	Probit		Bivariate Probit		
	LBW	**Marginal Effects**	**LBW**	**Marginal Effects**	**Prenatal Assault**
Prenatal Assault	0.267	0.037	0.879	0.183	
	[0.040]		[0.535]		
Lagged Incarceration Rate					-0.01
					[0.005]
Medicaid	0.002	0.000	0.002	0.000	0.169
	[0.002]		[0.002]		[0.018]
Black	0.263	0.036	0.262	0.035	0.25
	[0.013]		[0.013]		[0.094]
White	-0.093	-0.011	-0.094	-0.011	-0.122
	[0.012]		[0.012]		[0.089]
Single	0.026	0.003	0.026	0.003	0.209
	[0.003]		[0.003]		[0.021]
Over 35 years old	0.191	0.025	0.191	0.025	-0.05
	[0.003]		[0.003]		[0.033]
Teenage	0.096	0.011	0.096	0.011	0.01
	[0.003]		[0.003]		[0.020]
<HS	0.024	0.003	0.024	0.003	0.08
	[0.003]		[0.003]		[0.027]
HS graduate	0.021	0.002	0.021	0.002	0.042
	[0.003]		[0.003]		[0.024]
Male child	-0.049	-0.006	-0.049	-0.006	0.011
	[0.002]		[0.002]		[0.015]
Twin birth	1.649	0.459	1.649	0.459	0.012
	[0.004]		[0.004]		[0.046]
Father Black	0.049	0.006	0.049	0.006	0.114
	[0.009]		[0.009]		[0.067]
Father White	-0.037	-0.004	-0.036	-0.004	-0.067
	[0.008]		[0.008]		[0.064]
Father Hispanic	-0.042	-0.005	-0.041	-0.005	-0.106
	[0.008]		[0.008]		[0.062]
Father <HS	0.116	0.014	0.116	0.014	0.159
	[0.005]		[0.005]		[0.050]
Father HS	0.108	0.013	0.108	0.013	0.176
	[0.004]		[0.004]		[0.047]
Father some college	0.05	0.006	0.05	0.006	0.077
	[0.005]		[0.005]		[0.050]
Father information missing	0.232	0.031	0.232	0.031	0.373
	[0.005]		[0.006]		[0.051]
county unemployment rate	0.087	0.010	0.085	0.010	-1.008
	[0.043]		[0.043]		[0.357]
Real percapita income	-0.04	-0.005	-0.04	-0.005	-0.118
	[0.004]		[0.004]		[0.034]
Observations	4481243		4468224		4468224

Robust standard errors in brackets

Fig 1A: Female Hospitalizations for Assaults per 100,000

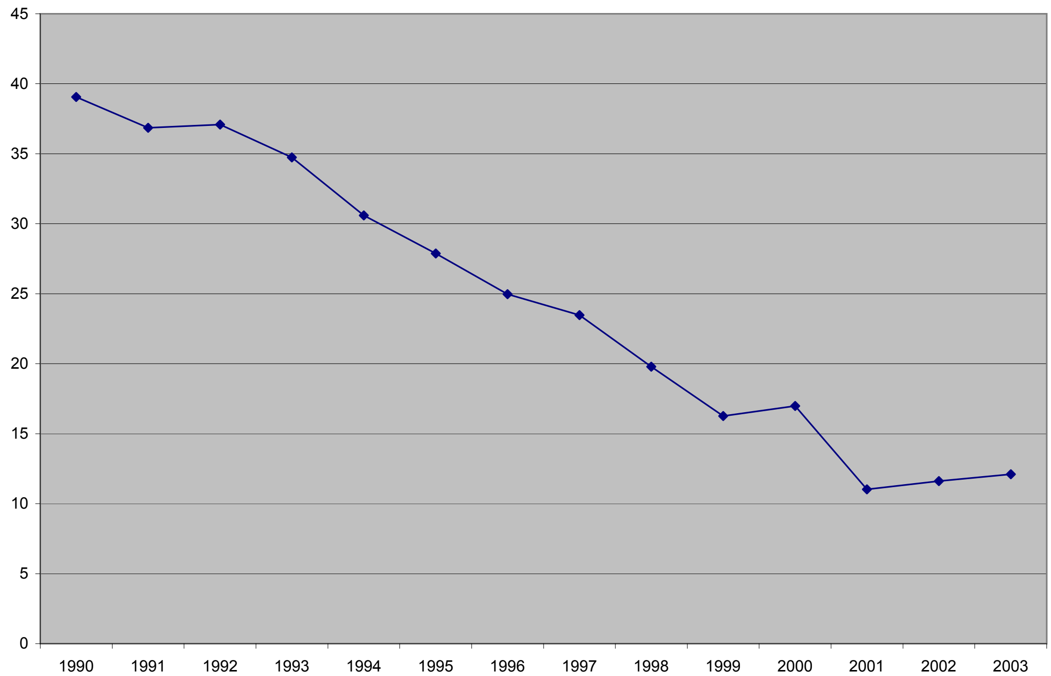


Fig 1B: Intimate Partner Homicides by Sex

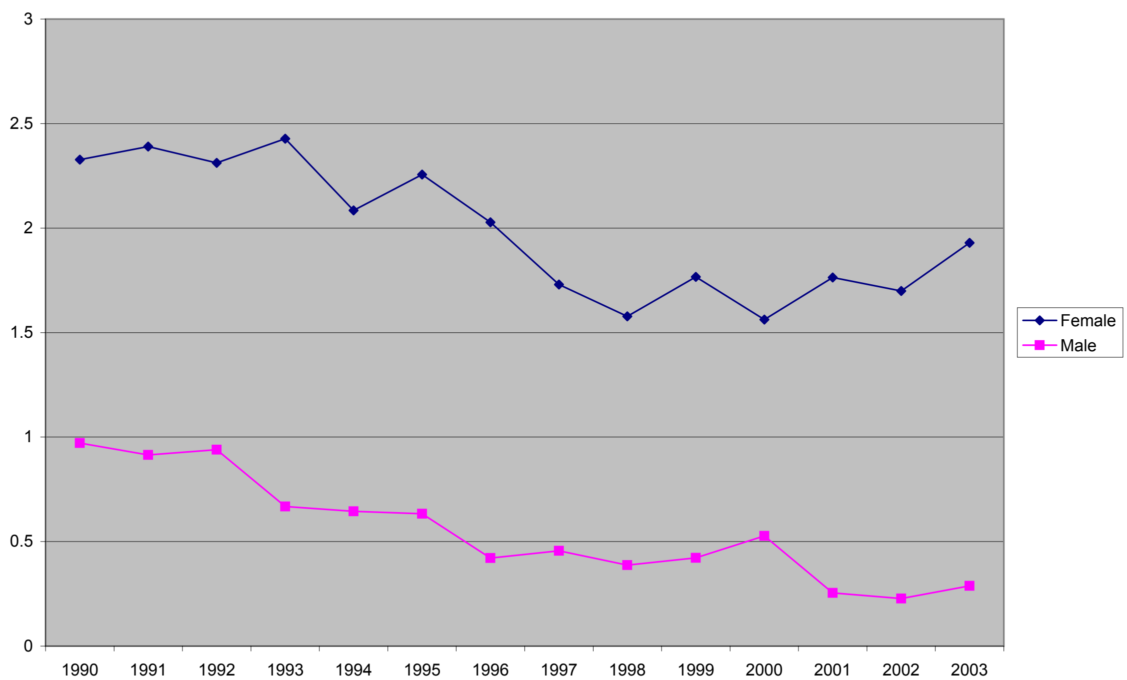


Fig 1C: Rates of Female Assault and Non Assault Injuries

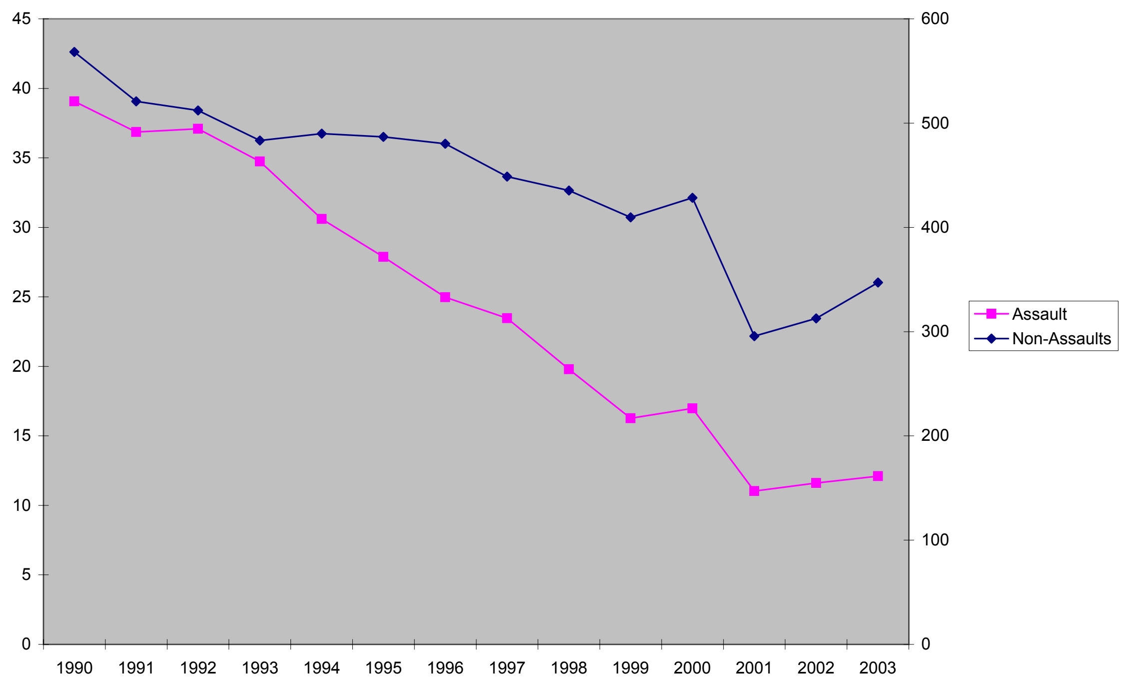


Fig 1D: Female Assaults as a Proportion of Total Injuries by Race

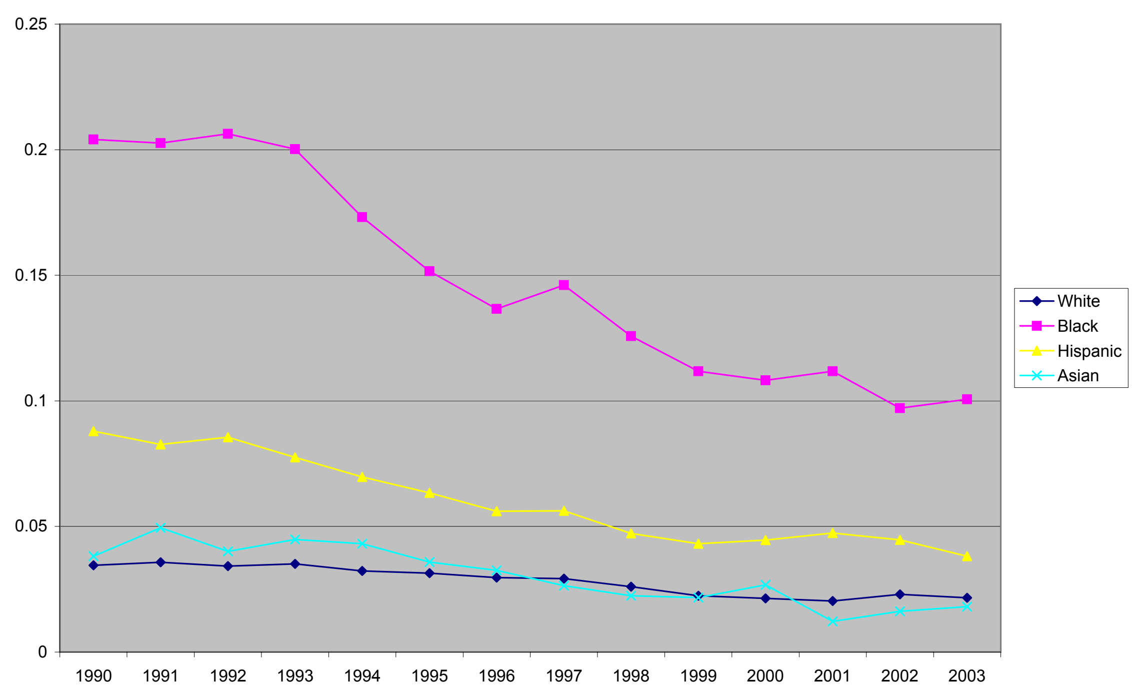


Fig 1E: Assaults as a Proportion of Total Injuries by Sex

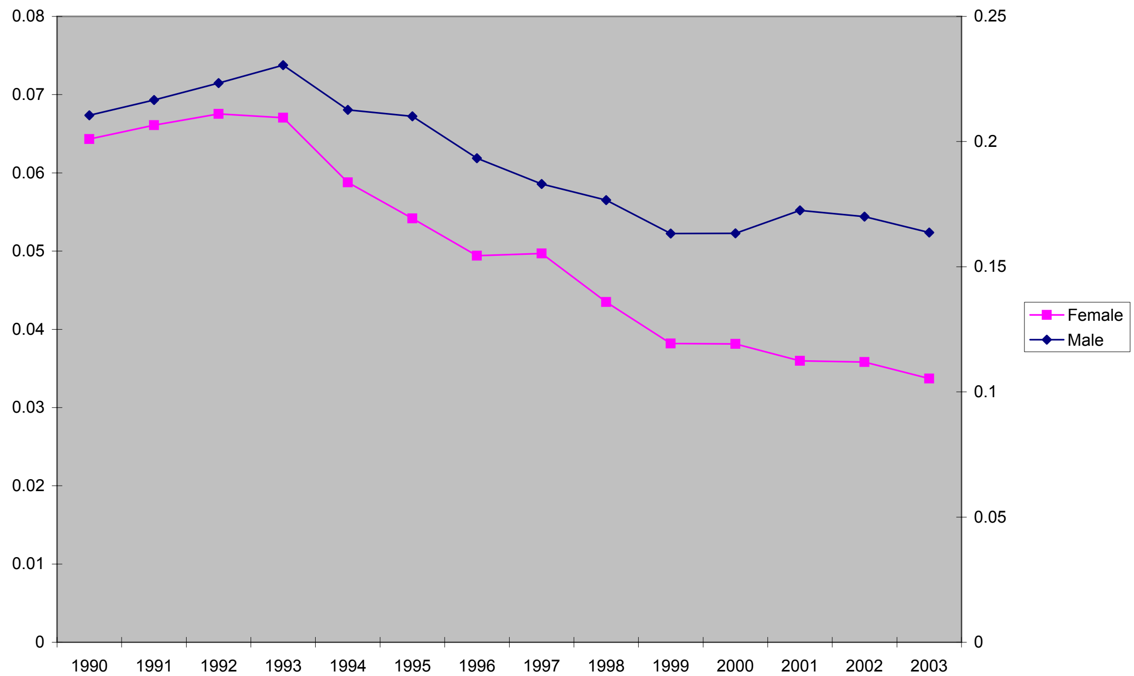


Fig 1F: Non-Intimate Homicides per 100,000

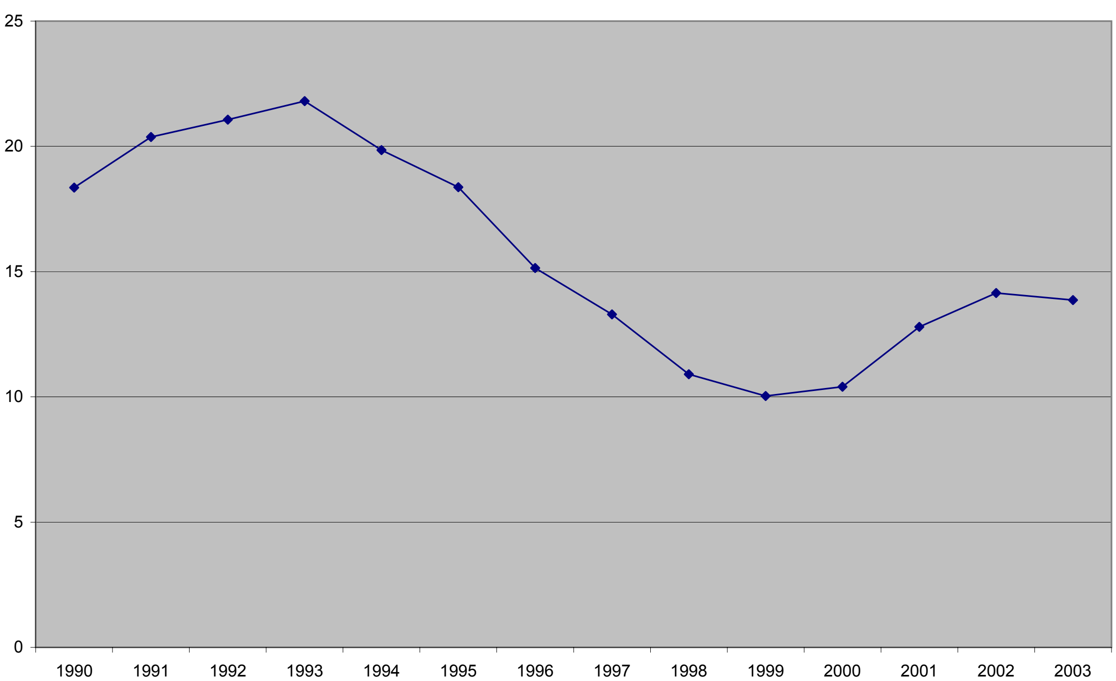


Figure 2 Propensity Score Estimates of Impact of Assault on LBW

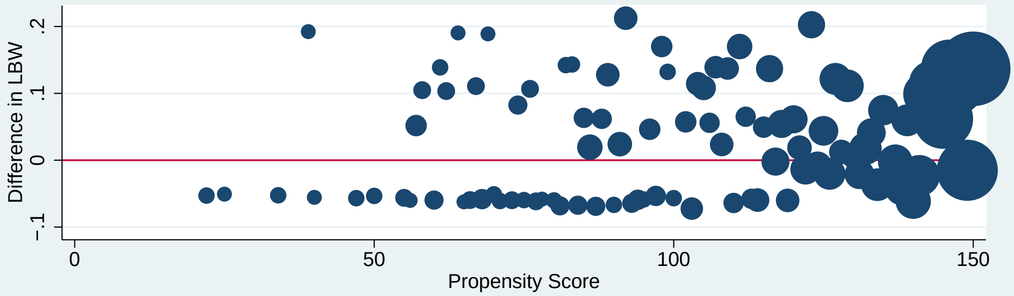
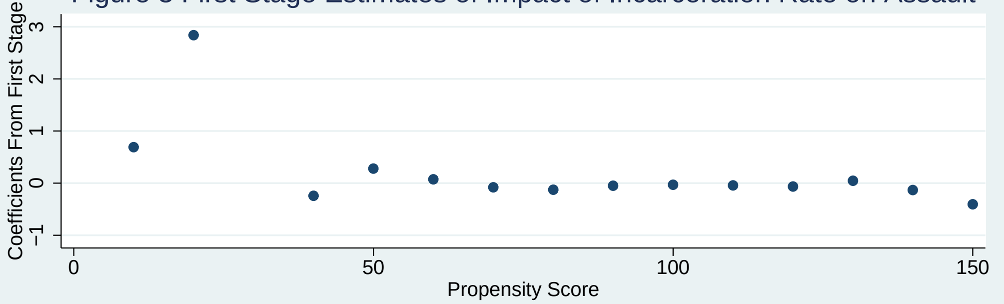


Figure 3 First Stage Estimates of Impact of Incarceration Rate on Assault



Appendix Table 1: Descriptive Statistics California Women's Health Survey (CWHS)

Sample Averages: CWHS 1998-2003

	Mean
Any domestic violence	0.053
Annual Personal Income of Women	29881
	[30886]
Female Labor Market Conditions (weekly wages)	233
	[60]
Age	39
White	0.546
Black	0.054
Hispanic	0.305
Asian	0.068
Other race	0.027
<HS	0.160
HS	0.238
<25 years old	0.140
25-30 years old	0.129
31-39 years old	0.267
40-49 years old	0.251
50-64 years old	0.213
Number of children <18	1.262
Single	0.180
Separated/Divorced	0.119
Cohabit	0.065
Married	0.669

Standard deviations in brackets below means

Probability of Violence and Average Income by Race, Education, Age and Marital Status

	Violence	Income
All	0.053	29881
White	0.043	32237
Black	0.072	26909
Hispanic	0.068	25837
Asian	0.035	34862
Other race	0.065	17281
<HS	0.072	24862
HS	0.063	23905
Some College	0.058	27354
College	0.029	39749
<25 years old	0.092	19306
25-30 years old	0.082	28767
31-39 years old	0.062	29798
40-49 years old	0.038	33174
50-64 years old	0.015	33705
Single	0.093	21098
Separated/Divorced	0.100	28973
Cohabit	0.090	27564
Married	0.035	32315

Appendix table 2: California Hospitalization & Employment Data Sample Means

	All	White	by Race Black	Hispanic
Wages, Employment				
Female Weekly Wage	217	241	226	187
Male Weekly Wage	260	279	246	251
Unemployment rate	0.073			
Ln(per capita income)	10.11			
Violence (per 100,000)				
Female Assaults	20.3	10.6	99.3	16.8
Female non-Assault Injuries	450			
Arrests	692	446	1700	841
Non Intimate Homicides	2.16	1.97	4.97	1.8
Car Crashes	113	125	140	88
Female Assaults - Older	7.1	5.6	23.6	6.3
Male Assaults	153	62	574	197
Male non-Assault Injuries	628			
Lagged incarceration rate	0.37	0.292	0.311	0.502

Appendix Figure 1:
Share of Women Across Industries - 1990 and 2000