Does government provision of healthcare explain the relationship between income inequality and low birthweight?

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This paper estimates the relationship between state and county income inequality and low birthweight (LBW) in the U.S. Specifically, it examines whether more unequal societies are also less healthy because such societies have lower investment in population health. The model includes an extensive list of community and individual controls along with community fixed-effects. I find that unequal states in fact have greater social investments, and absent these investments children born in such states would be more likely to be LBW. Using alternate measures of inequality reveals that income inequality in the upper-tail of the income distribution is not related to LBW; but inequality in the lower-tail of the income distribution is associated with increased LBW where the supply of healthcare mitigates the effect of income inequality. Consistent with prior findings, county income inequality is not significantly related to LBW.

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Introduction

That income and socioeconomic status protect health is well established (Newacheck 1994; Currie and Hyson 1999; Case, Lubotsky, and Paxson 2002; Currie and Stabile 2003). Moreover, it is generally agreed that the relationship between income and health, or the health production function is concave (Preston 1975; Rodgers 1979). What has been debated is whether income inequality has a detrimental effect on health. This paper examines one causal pathway through which state and county income inequality may operate upon infant health. I test whether income inequality is detrimental to health because unequal societies have lower expenditures on population health.

Low birthweight (LBW, birthweight < 2500 g) is an important outcome to study because LBW babies are more likely to be stunted and underweight (Osmani and Sen 2003), and increasing weight at birth increases the height, educational attainment, and earnings into adulthood (Behrman, Rosenzweig, and Taubman 1994; Currie and Hyson 1999; Behrman and Rosenzweig 2004). I examine the supply of health goods at the state (county) level as the causal link between state (county) income inequality and individual health because the provision of these goods is potentially modifiable through policy intervention.

The income inequality and infant health literature finds a positive correlation between income inequality and LBW (Kaplan et al. 1996; Lynch et al. 2001), pre-term birth (Huynh et al. 2005), and infant mortality (Rodgers 1979; Flegg 1982; Pampel and Pillai 1986; Waldmann 1992; Wennemo 1993; Kennedy, Kawachi, and Prothrow-Stith 1996; Fiscella and Franks 1997; Judge, Mulligan, and Benzeval 1998; Meara 1999; Mellor and Milyo 2001; Shi et al. 2004; Mayer and Sarin 2005). With the exception of a few studies that examine the relationship between income inequality and health at the individual level (Meara 1999; Mayer and Sarin 2005; Huynh et al. 2005), these studies use data aggregated at the country or state level.

At almost every level of aggregation (country, state, county, MSA, census tract, etc.) the negative correlation between inequality and health has been observed. However, very little is known about the precise mechanism by which income inequality poses a risk to individual health. The literature conjectures that social spending is one such mechanism, however, few studies have tested this hypothesis (Subramanian and Kawachi 2004; Mayer and Sarin 2005). The intuition behind this argument is that unequal societies are also unhealthier because the rich are unwilling to support policies or subsidize social goods that they are unlikely to benefit from or use. Empirical evidence on the government spending link between inequality and health is sparse and mixed. In a cross-country analysis, Pampel and Pillai (1986) find that controlling for government expenditures on healthcare renders the inequality effect on infant mortality statistically insignificant. Mayer and Sarin (2005) find that inequality is positively related with state spending on healthcare and reduces the probability of neonatal mortality. Although Shi et al (2004) do not test the government spending link, they find that the supply of physician care in the state is associated with lower LBW rate, and the inequality-infant health relationship is not completely explained by healthcare supply.

This paper contributes to the literature in the following ways: first, studies of income inequality and child health have largely focused on infant mortality. LBW as a child health outcome is relatively rare, and existing studies analyze data aggregated at the national or state level. To my knowledge, this is the first study to investigate the effect of income inequality on individual LBW. Mayer and Sarin (2005) find that public goods mediate the relationship between income inequality and neonatal mortality, and as the authors point out, this suggests that inequality may operate upon neonatal mortality first through its effect on birthweight. Second, social spending as a mechanism has been assessed at the state and country level, and this paper adds to the literature by analyzing the mechanism at the county level which is important because county governments are also responsible for making health financing decisions. Finally, prior studies of inequality at the sub-state level have suffered from limited statistical power (Subramanian and Kawachi 2004), and this paper uses data on the census of births to examine whether county inequality is related to infant health.

Income inequality is measured as the Gini coefficient (Atkinson 1970; Sen 1973; Foster and Sen 1999), because it is the most widely used measure and allows comparisons with other published works. The Gini has the property that any income transfer from the rich to the poor translates to a reduction in inequality. However, the inequality due to the rich getting richer is likely to have a different effect on public expenditures and health than inequality because the poor are getting poorer; this is not captured by the Gini. Following an anonymous referee's suggesting I the use the 90/50 and the 10/50 ratios to comment on this aspect of inequality.

Literature Review

Income inequality and health

A negative relationship between income inequality at the aggregate level, i.e. between geographic units, and average health in the communities is well established in the literature (Wilkinson 1992; Kaplan et al. 1996; Kennedy, Kawachi, and Prothrow-Stith 1996; Wagstaff and Doorslaer 2000). However, this relationship may simply be due to the concavity of the health production function[1] (Gravelle 1998; Subramanian and Kawachi 2004), and cannot be assumed to hold true at the individual level. Furthermore, income or other community characteristics may confound the income inequality effect. E.g., unequal communities may exhibit different lifestyle choices, different race/ethnicity, employment, occupation, and household composition characteristics due to factors correlated with but not caused by income inequality, compared to more equal communities. In the absence of adequate controls, income inequality may simply pick up the effect of omitted variables that are negatively associated with health (Hustead 1991; Levernier, Rickman and Partridge 1995; Phelps 1997; Partridge, Partridge, and Rickman 1998; Meara 1999; Bernard and Jensen 2000; Deaton and Lubotsky 2003; McLeod, Nonnemaker and Thiede-Call 2004).

In addition to community characteristics, several studies have estimated whether the income inequality effect persists after controlling for individual income and other demographic characteristics (Fiscella and Franks 1997, 2000; Daly et al. 1998; Kennedy et al.1998; Soobadeer and LeClere 1999; Meara 1999; Deaton and Paxson 2001; Mellor and Milyo 2001, 2002; Mayer and Sarin 2005). The results are mixed: for example, Mellor and Milyo (2002) find that the income inequality effect on self-rated health status disappears after controlling for household income, but Diez-Roux, Link and Northridge (2000) find that the income inequality effect on depressive symptoms and self-rated health status is attenuated but not eliminated with the inclusion of household income.

Another debate in the income inequality literature is -- what is the appropriate level of aggregation? Within the U.S., most studies have focused on the effect of state-level inequality on individual health (Daly et al. 1998; Kennedy et al. 1998; Blakely et al. 2000; Diez-Roux, Link, and Northridge 2000; Kahn et al. 2000; Lochner et al. 2001; Mellor and Milyo 2002; Subramanian, Kawachi, and Kennedy 2001; Subramanian and Kawachi 2003, 2004), but inequality at lower levels of aggregation, i.e., census tracts (Soobader and LeClere 1999), counties (Fiscella and Franks 1997) and metropolitan areas (Mellor and Milyo 2002, Blakely et al. 2002, Sturm and Gresenz 2003), have been considered. The results of these studies are mixed.

According to Wilkinson (1997), at higher levels of aggregation, i.e., within countries and states, income distribution is closely related to health, whereas income differences between states and countries do not affect health. Furthermore, he argues that when income inequality is measured in smaller areas, such as census tracts, it is not as important as income differences between these areas. Subramanian and Kawachi (2004) provide two explanations for why studies find state-level associations but only a weak association at lower levels of aggregation. First, one mechanism through which inequality affects health might operate at the state level since public expenditure for many components of healthcare, education and welfare are determined by taxing and spending decisions at the state level. Second, studies at lower levels of aggregation rely on smaller sample sizes thereby lacking the statistical power to find a significant correlation between inequality and health.

Income inequality and government spending, taxation and redistribution

The relationship between income inequality and public provision of health resources can be positive or negative. Theory predicts that greater income inequality may lead to greater demand for progressive taxation, assuming that the median voter is the decisive voter. When income inequality increases the median drops relative to the mean, and the median voter who is now poorer will demand greater pro-poor redistribution (Romer 1975; Roberts 1977; Meltzer and Richard 1981; Alesina and Perotti 1993; Alesina and Rodrik 1994). Milanovic (2000) finds that countries with greater pre-tax income inequality redistribute more to both the poor and the very poor. An alternate explanation for the positive association between income inequality and higher social spending is that as the poor get poorer, the rich get fearful of the poor (who may be driven to violence and crime) and choose to subsidize public goods and services in order to placate the poor (Piven and Cloward 1993).

Greater inequality may, on the other hand, be associated with lower social spending for two reasons. If the poor are disenfranchised and choose not to vote (Mayer and Sarin 2005); or if greater economic resources lead to more political influence, unlike the median voter theorem which assumes that political power is equally distributed (Benabou 1996; Rodriguez 1998), then the positive relationship between income inequality and redistribution may not hold. In a cross-country analysis Perotti (1996) finds that income inequality is positively associated with social security and welfare expenditure, but is not statistically significantly related to health, housing and education expenditures. Mayer (2003) in a state-level analysis finds that greater income inequality is not associated with an increase in AFDC spending, but is positively related to the number of AFDC recipients thus suggesting that the benefits per recipient declined.

Government spending and individual health

Population-wide interventions that improve women's access to care are expected to lower LBW, especially through care that is focused on smoking cessation, nutrition, and improving the general health of the mother (Alexander and Korenbrot 1995). Since the availability of healthcare is likely to lower the cost of health inputs (e.g. time cost) it is plausible that the public provision of healthcare will have a beneficial impact on child health. Government expenditures on health are associated with lower infant mortality (Judge, Mulligan, and Benzeval 1998), and the availability of government-subsidized family planning clinics has a large and negative effect on neonatal mortality in the U.S. (Grossman and Jacobowitz 1981). Changes in Medicaid eligibility have lowered infant mortality in the U.S. partly by increasing the probability of early prenatal care initiation and improved birthweights (Currie and Gruber 1996), and prenatal WIC participation is associated with greater adequacy of prenatal care and higher birthweight (Devaney, Bilheimer and Schore 1992).

Method

Figure 1 is a diagrammatic overview of the empirical strategy. The hypothesis is that inequality is correlated with worse health, and the mechanism by which inequality translates into poor health is that unequal societies have lower government provision of healthcare because the poor are either disenfranchised or lack political power to demand pro-poor transfers. As figure 1 shows, I estimate three equations to test this hypothesis: the first equation estimates the causal relationship between income inequality and government provision of healthcare, and I hypothesize that this relationship is negative; the second examines whether the public provision of healthcare is associated with better child health; and the final equation examines whether accounting for the mediating variable, government provision of healthcare, alters the relationship between income inequality and child health. I hypothesize that controlling for public spending on health will attenuate the negative effect of income inequality on birthweight.

The three estimation equations are as follows:

(1) $P_{st} = \beta_1 Q_{st} + \beta_2 X_{st} + \varepsilon_s + \nu_{st}$

Equation (1) estimates the relationship between income inequality, Q, in community s, at time t and the public provision of health care, P, in that community. β_1 represents the association between state (county) income inequality and supply of healthcare in the state (county), holding community characteristics, X, constant.

(2) $H_{is} = \delta_1 P_s + \delta_2 Z_{is} + \delta_3 X_s + \eta_{is}$

In equation (2), H indicates that child i in community s is LBW, δ_1 represents the relationship between public provision of healthcare and child health, holding constant maternal and child characteristics, Z, and community characteristics X.

(3) $H_{is} = \gamma_1 Q_s + \gamma_2 Z_{is} + \gamma_3 X_s + \gamma_4 P_s + \vartheta_s + \upsilon_{is}$

Equation (3) estimates the relationship between income inequality, Q, and LBW, H, controlling for mothers' characteristics, Z, and community characteristics, X. This model is estimated twice to compare the income inequality effect, γ_1, obtained from the specification that controls for public health provision, P, to that obtained from the specification without public health controls. If the introduction of P diminishes the coefficient on inequality, this will indicate that the income inequality and health relationship acts, at least in part, through the effect of inequality on public provision of healthcare. This strategy is similar to the one described in MacKinnon (2000)[2].

I hypothesize that (a) income inequality decreases public spending on health, i.e., β_1 is negative, (b) public provision of health goods share a positive relationship with health, i.e., δ_1 is positive, and (c) the inclusion of public health in the LBW equation (3) reduces the magnitude of the inequality effect, γ_1.

Ordinary Least Squares is used to estimate the equation when the outcome is continuous (e.g. state/county LBW rate), and probit estimation is used when the outcome is binary (LBW). Huber-White robust standard errors are estimated in equations 1-3. Further, because I use cross-sectional data from two time periods (1990 and 2000) I am able to include state (county) fixed-effects which eliminates any unobservable time-invariant state (county) characteristics correlated with state (county) income inequality or infant health. E.g. in equation (3) community culture (e.g. proclivity towards healthy behaviors) and political environment may be correlated with government expenditure on healthcare. The fixed-effects model exploits the within-state variation (over time) in income inequality to explain how changes in income inequality affect changes in infant health. Since individuals within a state (county) may be correlated, I re-estimate the models with standard errors clustered at the state (county) level to account for this within-cluster correlation. The results with and without clustered standard errors are virtually identical, and I present the latter set of results because estimating robust standard errors clustered at the state (county) level in a model with state (county) fixed effects may result in biased standard errors (Dow 2000).

In order to ensure that the results of this analysis are robust to community and individual controls, I enter each set of controls incrementally and note the effect controls have upon the explanatory variable of interest. Since the literature is divided on whether the income inequality effect is sensitive to these controls, the estimation strategy utilized in this analysis may help shed light on this ongoing debate. The final specification for each of the three equations includes the full set of state (county) and individual controls along with state (county) fixed-effects.

The specification in equations 1-3 has assumed a contemporaneous lag structure. While it seems sensible to assume that public provision of healthcare at the time of the pregnancy should have an effect on infant health, it is plausible that income inequality has a lagged effect on government expenditures. I model equations 1-3 under the assumption that inequality, public provision of healthcare, and birthweight are contemporaneously related for several reasons: first, theory provides no guidance on the appropriate length of the lag, and any lag structure that I impose would be necessarily arbitrary. Subramanian and Kawachi (2004) point out that the vast majority of income inequality studies make the same assumption. Nonetheless, since I have two years of data ten years apart I am able to test the effect of past inequality on current government expenditures and LBW. Results (not included) indicate that although past inequality is associated with present supply of healthcare, the relationship between ten years-lagged inequality and LBW is not robust to including community and individual controls. The second reason for assuming a contemporaneous lag structure is that current inequality is highly correlated with past inequality (Langer 1999; McLeod et al 2004), and hence serves as a good proxy for past inequality. Third, choosing the present specification makes the estimation more tractable since obtaining consistently measured data over multiple time periods from different sources is challenging, and utilizing data from two time periods enables fixed-effects estimation, a model with relatively conservative assumptions about the error terms.

Equation (1) assumes that income inequality has a causal effect on public healthcare provision but it is plausible that the reverse is true. I exploit the lag in the data and test whether public provision of health goods causes income inequality. This amounts to testing whether public health provision at time t-1 has a statistically significant effect on income inequality at time t, controlling for income inequality at t-1. The estimation equation is as follows:

(4) $Q_{st} = \alpha_1 Q_{s(t-1)} + \alpha_2 P_{s(t-1)} + \alpha_3 X_{st} + \omega_{st}$

I hypothesize that inequality at (t-1) will be highly correlated with inequality at time t, and including public health at (t-1) will not significantly alter α_1, the correlation between past and current inequality. Furthermore, a statistically insignificant coefficient on public health provision at (t-1), α_2, can be interpreted as evidence against the hypothesis that public health provision Granger causes income inequality. Admittedly, this test is not perfect because equation 1 assumes a contemporaneous relationship between income inequality and public health, whereas the Granger causality test assumes a lagged relationship. Nonetheless, the test serves to reduce some concern regarding the bias due to reverse causality.

The use of LBW as the dependent variable in this analysis brings up the question of selection because the probability of fetal survival may be correlated with the availability of healthcare (Pitt 2007). The fetal loss rate in the U.S. is approximately one in six pregnancies, including both early-term miscarriages and stillbirths, and this has remained stable through the 1990s (Ventura 2007). Any change in LBW due to the survival of fetuses that would have been lost without government provision of healthcare is likely to be small. Thus any potential bias in the government spending effect is expected to be small. In fact, fetuses that are saved as a result of increased government provision of healthcare are likely to be in the extreme left tail of the birthweight distribution. Thus, as a test I re-estimate the LBW equation (3) after eliminating the very LBW infants (birthweight < 1500 g) from my sample, and results using the restricted sample (available upon request) are virtually identical to those using the full sample of infants. This reduces the selection bias concern.

Data Description

The data for this analysis are drawn from several sources.

Individual data

Individual data are obtained from the 1991 and 2001 Natality Detail Files which provide information on the universe of live births in the U.S. The analysis is restricted to 1991 and 2001 because state and county measures of income inequality and community characteristics are available for those years. The Natality files compile birth certificates and provide information on birth outcomes, parental demographics, medical risk factors, prenatal care utilization and congenital abnormalities. Detailed information on mothers' place of residence at the time of birth is provided – 50 states and the District of Columbia are identified, and counties with population over 100,000, approximately 400, are identified. There are approximately 4,000,000 births in the U.S. every year, which yields an extraordinarily large sample and requires a great deal of computing resources. The analysis is thus restricted to a 15% random sample of singleton children born in 1991 and 2001, which yields a sample of 1,139,171 births. The county-level analysis is restricted to those counties with at least 50 observations. Although the Natality Detail files have sufficient observations to allow meaningful sub-state analysis, it provides no information on individual income. In order to overcome this limitation I include an extensive list of controls such as mothers' education, race, ethnicity, marital status, and several interactions of these key determinants of income.

Income inequality data

The 1990 and 2000 Decennial Censuses, specifically the 5% Integrated Public Use Microdata Series identifies all states and the District of Columbia, and provides information on individual income earned in 1989 and 1999, respectively. The individual income information is used to construct the state Gini coefficients, which capture income inequality in 1989 and 1999. The income data were calculated, topcoded and bottom coded as in Mayer and Sarin (2005).

To construct county Gini coefficients I use Census Summary Files, which provide the household distribution over 16 household income groups for 1989 and 1999, respectively, based on 1-in-6 samples from the Decennial Censuses long form. Following McLaughlin and Stokes (2002) and Huynh et al. (2005) I assign the midpoint of the income bracket to all households within a bracket and construct the county Gini coefficients for 1989 and 1999. This construction will potentially underestimate inequality within the county as the variation in household income in the Summary Files is underestimated because of the reliance on mid-points. Both the state and county Gini coefficients are standardized by subtracting year-specific mean and dividing the difference by year-specific standard deviation to aid in interpretation. The standardized Gini will have a mean 0 and variance 1, so a one unit increase in the standardized Gini is interpreted as a 1 standard deviation increase in the Gini.

Community characteristics

State and county characteristics are drawn from the 1990 and 2000 Census Summary Files. Community controls included in the model include median family income, racial and ethnic composition, education level (fraction 25 years or more with at least a college degree), and stability (percentage of owner-occupied homes that have been occupied for at least five years). The fraction of single female-headed households and fraction of single male-headed households in the states and counties control for family composition. Age controls (percent 65-74 years and percent greater than 75 years) are included because the hospital expenditure data used in this analysis do not distinguish between prenatal and other expenditures and a share of these expenditures may be made on behalf of elderly patients.

Healthcare provision data

The 2004 Bureau of Health Professions Area Resource File (ARF) yields the 1990 and 2000 per-capita number of hospitals beds and hospital expenditures by short-term general hospitals (STGH). STGH provide non-specialized care and the majority of patients stay for less than 30 days. STGH are an important source of prenatal care, especially for low-income, uninsured, and racial minority women (Perloff 1992). These data are available at the county level which I sum to the state level. It would be ideal to have data on state (county) owned hospitals, however, ARF, which is perhaps the most comprehensive source of public-use data on the availability of health resources, does not distinguish hospitals based on ownership. In 2005, the total number of hospitals in the U.S. was 5,759 of which 4,919 were community hospitals (non-federal STGH and specialty hospitals); approximately 60% of community hospitals were non-government nonprofit hospitals, 22% of the community hospitals were state or local government owned, and the remaining 18% were for-profit hospitals (AHA 2006). This suggests that the vast majority of STGH are likely to be government or non-government nonprofit hospitals. This may be an adequate proxy for government provision of healthcare because nonprofit hospitals, in exchange for tax subsidies, provide care for the poor. E.g. 80% of nonprofit hospitals in California provided uncompensated care, the total of which was twice the amount of subsidy received (DHHS 1997). Government hospitals are the largest providers of uncompensated care, but in some states nonprofit hospitals provide more uncompensated care than government hospitals (GAO 2005; CBO 2006). Additionally, regardless of ownership status, hospitals provide a wide range of community benefits, including parenting education, smoking cessation, and clinical services for indigent patients (GAO 2005).

Table 1 contains the means and standard deviations of the variables of interest, which shows that the probability of LBW in the 1991 and 2001 samples combined is 6%. Inequality rose at both state and county levels, a pattern that is in keeping with the overall inequality trend in the U.S. (Johnson, Smeeding, and Torrey 2005). The proportions of single female-headed households and college educated individuals also increased between 1991 and 2001. In the combined sample, the average number of STGH beds is 3.7 per thousand, and the average STGH expenditures per capita is approximately $1,082 ($1,161) at the state (county) level.

Results

Select regression coefficients are presented for the sake of brevity; the complete set of results is available upon request. Table 2 reveals the effect of state (county) income inequality on the two measures of healthcare in the state (county). Results show that income inequality is positively and statistically significantly correlated with the number of STGH beds per capita at the state and county levels. Including community fixed-effects and controlling for time-variant community characteristics including median income does not alter the magnitude or the statistical significance of this relationship. A one standard deviation increase in the Gini coefficient increases the number of beds per thousand by 0.265 (0.691), a change of 7.1% (19.6%) from the state (county) mean. On the other hand, a one standard deviation increase in the Gini coefficient decreases (increases) hospital expenditure, on average, by $70.65 ($35.64) per person in the state (county). This relationship is statistically significant at both the state and county levels. Thus, Table 2 shows that while the provision of beds increases as a result of greater inequality, hospital expenditures become less generous at the state level. These findings are consistent with Mayer and Sarin (2005) and Kaplan et al. (1996).

The result of the Granger causality test is reported in Table 3. It reveals that the 1990 Gini coefficient is highly (0.74) and statistically significantly correlated with the 2000 Gini coefficient. Including STGH beds and expenditures does not alter the correlation, and the coefficients on the two public health measures are not statistically significant. These results lend credence to the hypothesis that income inequality determines public health expenditures, and not vice versa.

Table 4 presents the relationship between health and STGH beds and expenditures. The top panel contains the results of the state-level analysis, and the bottom panel the results of the county-level analysis. Column 1 presents the correlation between each public health measure and LBW. The nature of the correlation is the same, regardless of whether government health provision is measured at the state or county level; number of STGH beds and expenditures per capita and LBW are positively associated, but most coefficients are statistically insignificant after including community and individual-level controls.

The inequality-LBW relationship is explored in Table 5. I estimate birthweight (grams) as an outcome as well. Results (not reported) are consistent with the findings of the LBW models discussed herewith. I begin by testing whether the LBW rate in the state (county) is positively related to inequality. Consistent with prior findings (Deaton and Lubotsky 2003; McLeod et al 2004), I find a positive and statistically significant association between state (county) inequality and state (county) LBW rate, but controlling for community characteristics such as median income and racial and ethnic composition makes the relationship insignificant. Columns 5-8 test whether state and county income inequality affects an infant's likelihood of being born LBW. Column 5 reveals that absent any controls, a unit increase in the standardized state Gini coefficient is associated with a 0.45 percentage point increase in the likelihood of LBW. Column 6 shows that the income inequality effect attenuates after state controls are included. Including state fixed-effects (column 7) strengthens the relationship between the Gini coefficient and LBW, which is in line with Wilkinson (1997) who argues that it is inequality within states that are related to health. Individual covariates weaken the inequality-LBW relationship, but including the public health measures, i.e. number of STGH beds and expenditure makes the income inequality effect larger. The income inequality coefficient reveals that a one standard deviation increase in the Gini coefficient leads to a 0.79 percentage point increase in the probability of LBW, which translates into 13.2% greater likelihood of LBW. Contrary to expectation, controlling for state health measures strengthens the income inequality effect on birthweight. This suggests that changes in the provision of health goods potentially dampen the adverse effect of state income inequality on LBW. This result is consistent with the findings related to neonatal and infant mortality (Mayer and Sarin 2005).

The lower panel presents the results of the county-level inequality analysis. The correlation between the standardized county Gini coefficient and LBW (column 5) is positive and statistically significant at the 0.01 level, however, including county characteristics as controls greatly reduces the magnitude of the correlation and renders it statistically insignificant. Including hospital beds and expenditure measures in the birthweight equation further diminishes the inequality coefficient. Thus Table 5 shows that while state inequality matters to individual health, county inequality does not. This is consistent with Subramanian and Kawachi (2004) and Wilkinson and Pickett (2006) who review the income inequality literature and note that inequality at the state or country level is closely related to health but at lower levels of aggregation (MSA and census tracts), inequality is weakly associated with health. Subramanian and Kawachi (2004) posit that perhaps the mechanism through which inequality affects health (e.g. public investment) operates at the state level. Wilkinson and Pickett (2006) explain that income inequality may be a proxy for social class differences in a society, and income inequality in smaller areas do not represent the degree of class stratification in the society as a whole. Thus, the seemingly contradictory results may reflect reality, i.e., it is state and not county inequality that is important in the determination of health. It is heartening that the results are consistent with prior findings but it remains possible that the difference in state and county results is due to data quality or collinearity. The county-level Gini coefficient is constructed using summarized income data which does not completely capture the effect of the full variation in income within counties on LBW. Alternatively, it is plausible that some of the community controls are collinear, and that there is greater collinearity in the county models than in the state models. E.g. the correlation between % black and median income is -0.16 at the state level and -0.31 at the county level; the correlation between % black and % single female headed households is 0.73 and 0.83 at the state and county levels, respectively.

Taken together, the results of the state-level analysis reveal that income inequality increases the provision of hospital beds but decreases the per capita hospital expenditures, and that the detrimental effect of inequality on the probability of LBW strengthens after controlling for the supply of healthcare. This suggests that health goods, as measured in this analysis, may reduce the adverse effect of inequality on birthweight. The county-level analysis reveals that the income inequality effect on health disappears after controlling for county characteristics; a similar result has been found elsewhere in the literature (Fiscella and Franks 1997; Blakely, Lochner and Kawachi 2002).

In order to test if these results are sensitive to the inequality measure used, I re-estimate equations 1 and 3 using the 90/50 ratio and 10/50 ratio. An increase in the 90/50 ratio signifies rising inequality because the rich are getting richer, and a decrease in the 10/50 ratio is interpreted as rising inequality because the poor are getting poorer. Results (available upon request) indicate that the rich getting richer within the state is associated with fewer STGH beds and expenditures in the state, but this form of inequality does not share a statistically significant relationship with LBW. On the other hand, the poor getting poorer is associated with greater provision of STGH beds and STGH expenditures in the state. Further, the relationship between the 10/50 ratio and LBW is statistically significant and the inclusion of government health variables in the model increases the magnitude of the inequality effect on LBW. This shows that a shift in the lower tail of the income distribution is accompanied by changes in government provision of health and LBW, and STGH beds and expenditure likely mediate the relationship between 10/50 ratio and infant health.

Conclusions

This paper estimates the relationship between income inequality (at both the state and county levels) and LBW. It explores whether hospital expenditure and the supply of hospital beds are the causal link between income inequality and LBW. At the state level I find that income inequality increases the number of hospital beds but decreases the average hospital expenditure. Contrary to expectation, however, controlling for the provision of healthcare does not mitigate the relationship between income inequality and LBW; in fact, the inequality coefficient is strengthened after including measures of healthcare provision in the model. Further, whilst the inequality – supply of healthcare – infant health relationship is sensitive to changes in inequality due to the poor getting poorer, the rich getting richer is not significantly associated with population health. At the county level, income inequality shares a statistically insignificant relationship with infant health, irrespective of the inequality measure used.

A limitation of this analysis is that hospital beds and hospital expenditures as measures of social spending or government expenditure on healthcare are one-dimensional. Although STGHs are an important source of prenatal care for low-income, uninsured, and racial minority women, using data on state/county Medicaid expenditures or government provision of prenatal care clinics would undoubtedly add richness to the current analysis. However, my best efforts failed to produce such data for 1990 and 2000 for either the states or the counties. Additionally, although these measures of health investment are obtained from the most comprehensive source of information on healthcare supply (ARF), they are not perfect because STGHs include federal, non-federal, and private hospitals, Thus, to be perfectly explicit, what this paper finds is that income inequality brings about a change in the supply of hospital beds and expenditures, but it cannot attribute this change to government intervention alone -- this change is most likely a combination of government intervention at many different levels of governance, market forces, and philanthropy.

This paper contributes to the literature in the following ways: First, there is considerable consensus that the relationship between inequality and health at the aggregate level disappears after controlling for the racial composition of the community (e.g. Deaton and Lubotsky 2003; McLeod et al 2004). Indeed that is what this paper confirms. But the relationship between state inequality and *individual* health, probability of LBW in this case, is robust to a vast array of state controls, including state median income and racial composition. Second, two recent reviews of the income inequality literature (Subramanian and Kawachi 2004; Wilkinson and Pickett 2006) have concluded that while health is related to inequality at higher levels of aggregation, the relationship is weak at lower levels. This paper tests the inequality-health relationship at two levels of aggregation (state and county) and confirms that this pattern holds. Finally, the main conclusion of this paper is that a change in the supply side of healthcare in response to a change in income inequality has the potential to alleviate the health risk posed by income inequality. It contributes to the income inequality literature by providing the following evidence -- contrary to expectation, on average, unequal societies have greater health investments and absent these investments children born in unequal societies are likely to be less healthy. These results are sensitive to where in the income distribution inequality is measured.

Since the negative effect of income inequality on child health persists after controlling for median income in the community, the message for policymakers is that the variation in the income distribution or income inequality is an important determinant of population well-being. Furthermore, while the mean of the income distribution is often used in making resource allocation decisions (e.g. Federal Medical Assistance Percentage for Medicaid[3]), this paper suggests that the variance of the income distribution is an additional yardstick for making government investment decisions.

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Table 1: Descriptive Statistics, 1991 and 2001 Natality Detail Files

Sample Means and (Standard Deviations in Parentheses)

	All		1991 Sample		2001 Sample	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Sample size	1,139,171		573,987		565,184	
Mean birthweight (grams)	3349.18	(576.29)	3358.87	(579.76)	3339.33	(572.58)
Percent LBW	6.00	(23.75)	5.96	(23.68)	6.04	(23.82)
Percent females	48.88	(49.99)	48.90	(49.99)	48.86	(49.99)
Mother's Characteristics						
Mean age at child birth	26.78	(6.02)	26.37	(5.83)	27.20	(6.18)
Mean years of education	12.63	(2.78)	12.43	(2.72)	12.83	(2.84)
Percent White	79.35	(40.48)	78.82	(40.86)	79.88	(40.09)
Percent Black	15.93	(36.60)	16.71	(37.31)	15.14	(35.85)
Percent other race	4.72	(21.21)	4.47	(20.65)	4.98	(21.75)
Percent Hispanic	18.47	(38.81)	15.46	(36.15)	21.54	(41.11)
Percent married	68.51	(46.45)	70.45	(45.63)	66.54	(47.18)
Percent foreign born	19.43	(39.57)	16.30	(36.94)	22.62	(41.84)
Percent MSA residents	81.16	(39.10)	79.97	(40.02)	82.38	(38.10)
State Characteristics						
Percent LBW	6.01	(1.42)	5.89	(1.53)	6.13	(1.30)
Average inequality (Gini coefficient)	43.03	(2.30)	42.36	(2.21)	43.69	(2.21)
Average median family income (1,000s 2000$)	45.63	(7.52)	43.45	(7.59)	47.82	(6.85)
Percent Black	10.71	(11.88)	10.61	(12.09)	10.80	(11.79)
Percent Hispanic	6.50	(8.20)	5.25	(7.38)	7.76	(8.83)
Percent single male headed households	5.10	(1.14)	4.37	(0.90)	5.83	(0.85)
Percent single female headed households	15.96	(4.46)	15.23	(4.31)	16.69	(4.52)
Percent 25+ with college degree or more	22.04	(4.89)	20.02	(4.16)	24.07	(4.75)
Percent homes occupied for 5+ years	50.92	(5.61)	50.73	(5.90)	51.12	(5.35)
Avg. no. short-term general hospital beds per-capita	3.73	(1.31)	4.23	(1.31)	3.24	(1.13)
Avg. short-term gen. hospital expenditures/capita (1,000s)	1081.70	(394.19)	835.96	(281.70)	1327.44	(333.67)
County Characteristics						
Percent LBW	6.04	(1.07)	5.98	(1.12)	6.10	(1.02)
Average inequality (Gini coefficient)	40.00	(3.18)	39.73	(3.26)	40.24	(3.09)
Average median family income/1,000 (2000 $)	49.12	(10.50)	46.69	(9.82)	51.29	(10.63)
Percent Black	11.79	(12.98)	11.81	(12.77)	11.78	(13.17)
Percent Hispanic	8.35	(12.64)	6.53	(10.88)	9.98	(13.83)
Percent single male headed households	5.06	(1.25)	4.31	(0.95)	5.73	(1.09)
Percent single female headed households	16.60	(5.30)	16.04	(5.24)	17.10	(5.31)
Percent 25+ with college degree or more	23.69	(8.45)	21.42	(7.46)	25.72	(8.77)
Percent homes occupied for 5+ years	49.24	(7.23)	48.80	(7.65)	49.63	(6.83)
Avg. no. short-term general hospital beds per-capita	3.53	(2.27)	4.09	(2.43)	3.03	(2.00)
Avg. short-term gen. hospital expenditures/capita (1,000s)	1160.76	(810.88)	903.15	(553.93)	1387.70	(926.19)

Notes:

State Gini coefficient constructed from the 1990 and 200 5% IPUMS; county Gini coefficient constructed from 1990 and 2000 Decennial Censuses; state and county demographic, education, household characteristics obtained from the 1990 and 2000 Dencennial Cen

Table 2: Relationship Between Income Inequality (Gini Coefficient) and Supply of Health Goods

OLS Coefficients on Standardized Gini Coefficient and (Standard Errors in Parentheses)

	(1)	(2)
Dependent Variables:		
State-Level Public Goods		
No. short-term general hospital beds per-thousand population	0.228**	0.265**
	(0.001)	(0.001)
Short-term gen. hospital expenditures/capita	52.937***	-70.654***
	(0.314)	(0.391)
County-Level Public Goods		
No. short-term general hospital beds per-thousand population	1.001***	0.691***
	(0.002)	(0.007)
Short-term gen. hospital expenditures/capita	370.393***	35.641***
	(0.868)	(3.121)

Notes:

Gini Coefficient measures income inequality.

Gini standardized by subtracting year-specific mean and dividing by year-specific standard deviation.

Each pair of OLS coefficient and s.e. derived from a separate regression

Column (1) contains OLS coefficients on Gini coefficient, with year fixed-effects

Column (2) contains OLS coefficients on Gini coefficient, controlling for community characteristics -- Median family income, % Black, % Hispanic, % Single male headed hh, % Single female headed hh, % 25+ w/ college degree or more, % homes occupied 5+ years, % 65-74 years, % 75+, and year and community fixed-effects

*(**)(***) represent statistical significance at the 0.1(0.05)(0.01) levels, respectively

Table 3: Granger Causality Test: Relationship Between Lagged Public Health Goods and Current Gini Coefficient, Controlling for Laggest Gini

OLS coefficients (and standard errors in parentheses)

Dependent Variable: Standardized Gini Coefficient, 2000

	Step 1	Step 2
Standardized Gini Coefficient, 1990	0.744 ***	0.765 ***
	0.075	0.076
No. short-term general hospital beds per-thousand	-	-0.053
		0.059
Short-term gen. hospital expenditures/capita	-	-0.00021
		0.0004

Notes:

Gini Coefficient measures income inequality.

Gini standardized by subtracting year-specific mean and dividing by year-specific standard deviation.

Step 1 and 2 control for community characteristics -- Median family income, % Black, % Hispanic, % Single male headed hh, % Single female headed hh, % 25+ w/ college degree or more, % homes occupied 5+ years, % 65-74 years, % 75+, and year and community f

*(**)(***) represent statistical significance at the 0.1(0.05)(0.01) levels, respectively

Table 4: Relationship Between Supply of Healthcare (State and County Levels) and Probability of Low Birthweight, 1991 and 2001 Natality Detail Files

Probit Marginal Effects and (Standard Errors in Parentheses)

	(1)	(2)	(3)	(4)
State-Level Public Goods				
No. short-term general hospital beds per-thousand population	0.00655***	0.00183***	0.00470**	0.00355*
	(0.00028)	(0.00051)	(0.00216)	(0.00209)
Short-term gen. hospital expenditures/capita	0.00002***	0.000002	0.00001	0.00001
	(0.000001)	(0.000002)	(0.00001)	(0.00001)
Sample size	1,139,171	1,139,171	1,139,171	1,139,171
County-Level Public Goods				
No. short-term general hospital beds per-thousand population	0.00438***	0.00094***	0.00106**	0.00047
	(0.00012)	(0.00018)	(0.00042)	(0.00041)
Short-term gen. hospital expenditures/capita	0.00001***	0.000002***	0.0000006	0.0000006
	(0.0000003)	(0.0000005)	(0.000001)	(0.000001)
Sample size	831,191	831,191	831,191	831,191
Community Controls	No	Yes	Yes	Yes
Individual Controls	No	No	No	Yes
Community Fixed-effects	No	No	Yes	Yes

Notes:

Gini Coefficient measures income inequality.

Gini standardized by subtracting year-specific mean and dividing by year-specific standard deviation.

Low Birthweight=1 if birthweight <-2500 g; 0 otherwise.

Community controls include: Median family income, % Black, % Hispanic, % Single male headed hh, % Single female headed hh, % 25+ w/ college degree or more, % homes occupied 5+ years, % 65-74 years, % 75+, and year fixed-effects

Individual controls include: child sex, mother's age, years of education, race and ethnicity, marital status, foreign born status and MSA residence status

*(**)(***) represent statistical significance at the 0.1(0.05)(0.01) levels, respectively

Table 5: Relationship Between Income Inequality and LBW -- With and Without Supply of Healthcare, 1991 and 2001 Natality Detail Files

	Low Birthweight Rate parentheses)				Probability of Low Birthweight Probit Marginal Effects and (Standard Errors in parentheses)				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
State-Level Public Goods									
Standardized Gini coefficient	0.007***	0.00002	0.005	0.004	0.0045***	0.0011**	0.0075***	0.0063***	0.0079***
	(0.001)	(0.001)	(0.005)	(0.005)	(0.0002)	(0.0005)	(0.002)	(0.002)	(0.0024)
No. short-term general hospital beds per-thousand population				0.003					-0.002
				(0.004)					(0.003)
Short-term gen. hospital expenditures/capita				0.00001					0.00001*
				(0.00001)					(0.00001)
Sample size	102	102	102	102	1,139,171	1,139,171	1,139,171	1,139,171	1,139,171
County-Level Public Goods									
Standardized Gini coefficient	0.0084***	0.00120**	-0.0037	-0.0034	0.0088***	0.0013**	0.0032	0.0015	0.0014
	(0.0007)	(0.0010)	(0.0033)	(0.0034)	(0.0026)	(0.0006)	(0.002)	(0.002)	(0.002)
No. short-term general hospital beds per-thousand population				0.00028					0.00009
				(0.0010)					(0.0005)
Short-term gen. hospital expenditures/capita				-0.000001					0.0000003
				(0.000002)					(0.000001)
Sample size	706	706	706	706	831,191	831,191	831,191	831,191	831,191
Community Controls	No	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes
Individual Controls	No	No	No	No	No	No	No	Yes	Yes
Community Fixed-effects	No	No	Yes	Yes	No	No	Yes	Yes	Yes

Notes:

Gini Coefficient measures income inequality.

Gini standardized by subtracting year-specific mean and dividing by year-specific standard deviation.

Low Birthweight=1 if birthweight <=2500 g; 0 otherwise.

Community controls include: Median family income, % Black, % Hispanic, % Single male headed hh, % Single female headed hh, % 25+ w/ college degree or more, % homes occupied 5+ years, % 65-74 years, % 75+, and year fixed-effects

status

*(**)(***) represent statistical significance at the 0.1(0.05)(0.01) levels, respectively

Figure 1: How government healthcare expenditure mediates the negative relationship between income inequality and birthweight: hypothesized relationships and theoretical context

Hypothesis 1

Independent variable: Income inequality, Q → Mediator: Government expenditure on healthcare, P

(-)

Rich don't support pro-poor transfers
Poor disenfranchised and don't vote

Hypothesis 2

Mediator: Government expenditure on healthcare, → Dependent variable: Birthweight, H

(+)

Short-term General hospitals are an important source of prenatal care, especially for low-income mothers

Hypothesis 3

Independent variable: Income inequality → Dependent variable: Birthweight

(-)

Magnitude decreases after including mediating variable (gov't expenditure on healthcare)

Mediator: Government expenditure on healthcare

[1] Assuming a concave relationship between individual income and individual health, if $x is transferred from the richest person to the poorest person (thus reducing inequality), then the poor person's health improvement will be more than the rich person's health deterioration, thereby leading to an improvement in the average health of society.

[2] Equation 1 in this analysis is equivalent to MacKinnon's step 2, and estimating equation 3 with and without the mediator variable is tantamount to estimating MacKinnon's steps 1 and 3.
[3] The federal government's share of a state's Medicaid costs is inversely proportional to the state per-capita income relative to the national average (Scott 2005).