Institute for Research on Poverty
Discussion Paper no. 1212-00

Welfare's Children

Michael Wiseman
Institute for Research on Poverty
University of Wisconsin–Madison

and

National Opinion Research Center
Washington, DC
Email: MichaelLWiseman@cs.com

September 2000

Research for this study was supported by the Urban Institute and by the California Department of Social Services through contract with the Survey Research Center of the University of California at Berkeley. Carol Wiseman performed the computations. The paper includes some material from Wiseman (2000) that is reproduced verbatim without citation.

Abstract

States with family cap public assistance policies deny or reduce additional welfare benefits to mothers who conceive and give birth to additional children while they are receiving aid. By 1999, 22 states had family cap policies in place. Little is known about the incidence of conceptions leading to births among women on welfare or the effect of family cap policies on the likelihood of such events. This paper reports estimates of the number and cost implications of infants conceived by mothers receiving assistance in California over the period 1988–1995. The estimates are constructed using a longitudinal analytic database derived from administrative records. The results indicate that such births are common, but their incidence is declining. In 1988 about 5.2 percent of the AFDC-FG (mostly single-parent) California cases experienced a birth of a child conceived while the case was open; by 1995 this had declined to 4.8 percent. Over the same interval this aggregate rate fell from 8.0 to 6.7 percent for (two-parent) families in AFDC-UP. Although the one-year number of such births (62,000 in 1995) is small in comparison with total numbers of children in AFDC families, such children accumulate over time. Thirty percent of children under age 9 in California families receiving AFDC benefits in February 1996 were conceived while their cases were open; benefits paid on their behalf amounted to roughly 7 percent of total state outlays. The incidence of such births varies by race, but the downward trend is common. California actually implemented a family cap in 1997, but the state's plan is unusual in that adults can escape the consequences by leaving welfare for two months following the onset of pregnancy. Because administration of TANF in California has been radically decentralized, it is not clear whether counties are in fact applying the cap or, if so, whether parents are taking advantage of the two-month exemption.

Welfare's Children

States with family cap public assistance policies deny or reduce incremental welfare benefits to mothers who conceive and give birth to additional children while they are receiving aid. Such policies are both controversial and common. Between January 1992 and the passage of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) in August 1996, 19 states sought and received approval (waivers) from the U.S. Department of Health and Human Services to experiment with family caps of one variety or another (U.S. Department of Health and Human Services, 1997). The same 19 states included some sort of cap in the Temporary Assistance for Needy Families (TANF) plan formulated to meet the requirements of PRWORA. By July 1999 the number of states with family cap policies in place had increased to 22 (Rowe, 2000). California, the state with the largest dependent population in the 1990s, applied for and received a federal waiver for a family cap in 1993, but the policy was not implemented until 1997. This paper sets the stage for evaluating the consequences of the cap both in California and elsewhere by studying childbearing among California's welfare recipients in the pre-PRWORA era.

Table 1 lists states with family caps in 1999. At that time, these states accounted for about half of the national TANF caseload and about half of all poor families with children. It is therefore reasonable to say that about half of all women at risk of bearing children potentially eligible for cash assistance were in 1999 making decisions in a regulatory environment that included the cap. Though circumstances vary, in general state family cap policies are motivated by public belief that welfare use is in part the result of irresponsible behavior and that subsequent conceptions among recipients are more of the same. The caps are generally part of a larger collection of policies intended to reduce long-term welfare use and to promote movement by the welfare-dependent from public assistance to greater self-sufficiency. For example, Wisconsin's benefit cap demonstration project was expected, according to state literature, to "reinforce work, responsibility, and family" (Wisconsin Governor's Office, 1996).

TABLE 1
State Family Cap Provisions

State	Family Cap	Implementation Date	Child Excluded if Born More than X Months after Case Opening	Increase in Cash Benefit (and Special Provisions)	Capped Child Eligible if Case Closed X Months[a]
Arizona	Yes	Nov-95	?[b]	None (disregard)[c]	60
Arkansas	Yes	Jul-97	9	None	?
California	Yes	Aug-97	10	None	24
Connecticut	Yes	Jan-96	10	$50	?
Delaware	Yes	Oct-95	10	None	?
Florida	Yes	Oct-96	?	Half of standard increment	?
Georgia	Yes	Jan-94	10	None	?
Idaho	Implicit[d]	Jul-97	—	None (disregard)	—
Illinois	Yes	Jan-96	10	None	9
Indiana	Yes	May-95	10	None	Always capped
Maryland	Yes	Apr-96	10	None (third party payment)[e]	?
Massachusetts	Yes	Oct-95	10	None (disregard)	Always capped
Mississippi	Yes	Nov-95	10	None	Always capped
Nebraska	Yes	Nov-95	10	None	6
New Jersey	Yes	Oct-92	10	None (disregard)	12
North Carolina	Yes	Jul-96	9	None	?
North Dakota	Yes	Jul-98	8	None	12
Oklahoma	Yes	Nov-97	10	None (voucher)[f]	?
South Carolina	Yes	Jan-97	10	None (voucher)	?
Tennessee	Yes	Sep-96	10	None	1
Virginia	Yes	Jul-95	10	None	Always capped
Wisconsin	Implicit	Jan-96	—	None	—

Source: The Urban Institute Welfare Rules Database. For details on special provisions, see Rowe (2000).
[a]Indicates the number of months a unit must remain off of assistance in order to regain eligibility for a previously capped child.
[b]? indicates provision not apparent in state operations manuals.
[c](disregard) indicates state increased amount of earned income disregarded in benefit computation when family is subject to a cap.
[d]States with implicit caps have a benefit that is fixed regardless of family size.
[e]Increment in benefit for child subject to cap goes to a third party for use on child's behalf.
[g]Standard incremental child benefit paid in vouchers for child-related expenses.

Family cap initiatives offer an interesting case for studying the interaction between research and policymaking. Despite the political attractiveness of caps, there is little empirical support for expecting them to do much beyond reducing costs. By far the dominant conclusion of the literature on welfare effects on fertility is that such influences, though present, are small and uncertain. In a recent survey of this literature, Robert Moffitt summarizes over 40 studies by stating that “a neutral weighing of the evidence still leads to the conclusion that welfare has incentive effects on marriage and fertility, but the uncertainty introduced by the disparities in research findings weakens the strength of that conclusion” (Moffitt, 1998, p. 75).

Family caps ostensibly act only on the likelihood that a woman receiving assistance on behalf of one or more children already born will become pregnant and bear another. Benefit effects on this decision have been less widely studied than effects on fertility in general. Using data from the National Longitudinal Survey of Youth covering the period 1979–1988, Acs (1996) finds no effect of Aid to Families with Dependent Children (AFDC) benefit levels or increments on the likelihood that young women will have additional children, whether or not they are on welfare. Using data from the Survey of Income and Program Participation (SIPP) for the period 1989–1992, Fairlie and London (1997) find some indication of positive correlation between incremental AFDC benefits and the likelihood that an AFDC recipient will have an additional child. However, they discover an even stronger association between incremental benefits and the fertility of various *non*recipient groups. The result is a muddle; it is not clear at all what is leading to the partial correlation between state policy and welfare fertility in the SIPP data.

Could it be done, classical experimentation might better identify any effect of financial incentives on welfare fertility. This is the motivation of the federal government in promoting random assignment as a method for evaluating the effects of family cap and other policies. Since the Acs and Fairlie/London papers were published, two evaluations of state family cap provisions based on random assignment have appeared. Turturro, Benda, and Turney (1997) report no evidence of effects on fertility of the waiver-based experiment that Arkansas began in 1994. Camasso, Jagannathan, and Killingsworth (1998a, 1998b) found that for both ongoing and newly opened cases, introduction of a cap in New Jersey in 1993 reduced birth rates among families subject to the new provision relative to birth rates within a control group exempted from the provision. Unfortunately, both the Arkansas and New Jersey evaluations suffer from serious problems of implementation and analysis (Loury, 2000; Rossi, 2000). While the New Jersey evaluation appears both in terms of implementation and methodology to be superior to that for Arkansas, Rossi's conclusion that "the deficiencies [of the New Jersey evaluation] . . . are serious enough to cast strong doubts on the validity of the findings" leaves the muddle unresolved.

The Acs, Fairlie/London, Arkansas, and New Jersey studies are based on household data, in some cases drawn from administrative sources. Two other studies address cap effects using more aggregate analysis. The White House Council of Economic Advisers (CEA) has looked at the effect of state family caps on growth of welfare caseloads (CEA, 1999). As the Council acknowledges, it does not seem that imposition of a family cap should reduce welfare caseloads. The women at risk of additional pregnancies and therefore subject to the cap are already receiving assistance. Therefore even if a family cap is effective in influencing subsequent fertility, such a policy, when introduced , would not for most cases alter immediate welfare eligibility and therefore should not affect caseloads. Despite this lack of mechanical connection between family caps and the caseload, the CEA's caseload model links the advent of state experimentation with such policies over the period 1992–1996 to *increasing* rates of growth of welfare use (CEA, 1999). On the other hand, using data from 1984 through 1996, Horvath and Peters (1999) find that state experiments with family caps in AFDC reduced the share of births that were nonmarital. In contrast to the CEA's results, the Horvath and Peters outcome would be expected to reduce caseloads, especially if the effect occurs for first births. Just what is going on in the partial correlations uncovered in these studies is still far from clear—again, a muddle. At minimum, the implication is that researchers should be cautious.

Beyond the uncertainty about causal connection between financial incentives and welfare childbearing, there is an issue of numbers. Judged from textbooks, every introductory policy analysis course features some review of steps to be undertaken in evaluating alternatives for tackling problems that are thought appropriate for government attention. Invariably, the template calls for beginning the analysis with assessment of the problem. However, none of the waiver requests made by states prior to PRWORA to permit imposition of family caps included data on the incidence of conceptions by mothers while they were receiving welfare benefits (hereafter called welfare conceptions).

At least two reasons might be cited for this rare departure from the rational policy paradigm. One is found at the federal level, the other at the state level. The federal agency responsible for granting permission for state welfare experiments, the Administration for Children and Families within the U.S. Department of Health and Human Services, focused its attention principally upon assuring that the innovations were evaluated by random assignment and not upon procedures followed by states in selecting policies to pursue. On the states' side, implementation of a family cap calls for linking newborns to the circumstances of their parent(s) at the time of conception. Until very recently, most state case management systems were designed primarily for dealing with *current* transactions—assessment of eligibility, calculation of payments, allocation of current services, and the like (Wiseman, 1999). Even when data maintained for current transactions was in principle adequate for discriminating between births subject to the cap and births exempted from it, the programming task necessary to develop retrospective data on mothers' circumstances in the month of conception is daunting. Since the federal government did not press, the states avoided the effort.

Are welfare conceptions sufficiently common to justify so much attention? To lay a foundation for analysis of the consequences of the new wave of caps, this paper steps back and investigates what will be called the numbers question: How common are welfare conceptions, and how has childbearing among welfare recipients changed over time? A new and richer administrative data set is used to count the incidence of welfare conceptions and the number of children on welfare who might be subject to a family cap in the nation's largest state, California. The data cover the period 1988–1996 and therefore allow assessment of changes in the quantity and incidence of welfare conceptions over time. The calculation sets a baseline for analysis of trends after implementation of the state's new, post-block-grant initiative, CalWORKS.

In addition to providing exceptionally detailed information on welfare conceptions, this study provides insight into both the utility of administrative data for policy research and problems that can arise in such efforts. Administrative data have long been used by evaluators for analyzing the effects of welfare-to-work programs. By focusing on the welfare/fertility connection over time in a particular state, this paper breaks new ground. The recent report of the Joint Center for Poverty Research on research uses of administrative data encourages the development of analytic data from administrative sources for research purposes (Advisory Panel, 1998; see also Winn and Lennon, 2000). The present effort illustrates both the promise of such efforts and the problems commonly encountered by researchers struggling to employ administrative data for social science.

The next section sets the stage by reviewing California's pre-PRWORA welfare system and trends in welfare use.

BACKGROUND: WELFARE IN CALIFORNIA

Eligibility and Benefits

As elsewhere, California families with children qualified for AFDC if their incomes adjusted for certain expenses fell below the state's "standard of need" and (a) one parent was absent from the household or disabled or (b) the family's "principal earner" was involuntarily working less than 100 hours per month. Families qualified on basis (a) were included in AFDC-FG (for “family groups”); families qualified on basis (b) were included in AFDC-UP (for “unemployed parent”). In California the unemployed parent program was normally labeled AFDC-U, but for this paper the nationally more common UP designation is used.

Through 1989, California’s welfare benefit was increased each year to keep up with inflation. This adjustment was eliminated beginning in 1990, and the nominal benefit itself was reduced in 1991, 1992, and 1993. Figure 1 reports the results of both changes, showing real AFDC benefit levels for a family of two over this interval and beyond, along with the increment in benefits associated with an increase in the family’s size by one.[1]

Between the peak in 1986–89 and the trough in 1998, the value of the combined AFDC and food stamp benefit for a family of two fell by 22 percent in California. The benefit increase brought by additional children fell by less—19 percent. After 1997 the legislature moved to increase benefits (and to add a cap), but even for cases not subject to the cap, the increase in the marginal benefit associated with a move from one to two children (2.7 percent) was less than the gain in the level of benefit for a two-person family (3.6 percent). While California reduced its welfare payment, the welfare eligibility standard, called the Minimum Basic Standard of Adequate Care, continued to increase. The consequence was that families eligible for AFDC in 1988 would also have been eligible had they experienced similar circumstances in 1996, but the cash benefit was lower.

The California Welfare Caseload

The decade from 1985 through 1995 was one of exceptional caseload growth. Figure 2 shows the aggregate AFDC caseload and the two components, AFDC-FG and AFDC-UP. Between 1988 and 1995 the AFDC caseload in the state increased by 54 percent. By 1995, 21 percent of the state’s children were

---

[1]These series are deflated by the Consumer Price Index less medical costs. AFDC recipients are eligible for Medi-Cal, the state’s Medicaid program. Medi-Cal services did not change significantly over this interval.

FIGURE 1
AFDC/TANF and Food Stamp Benefits
for California Family of Two, with Birth Increment

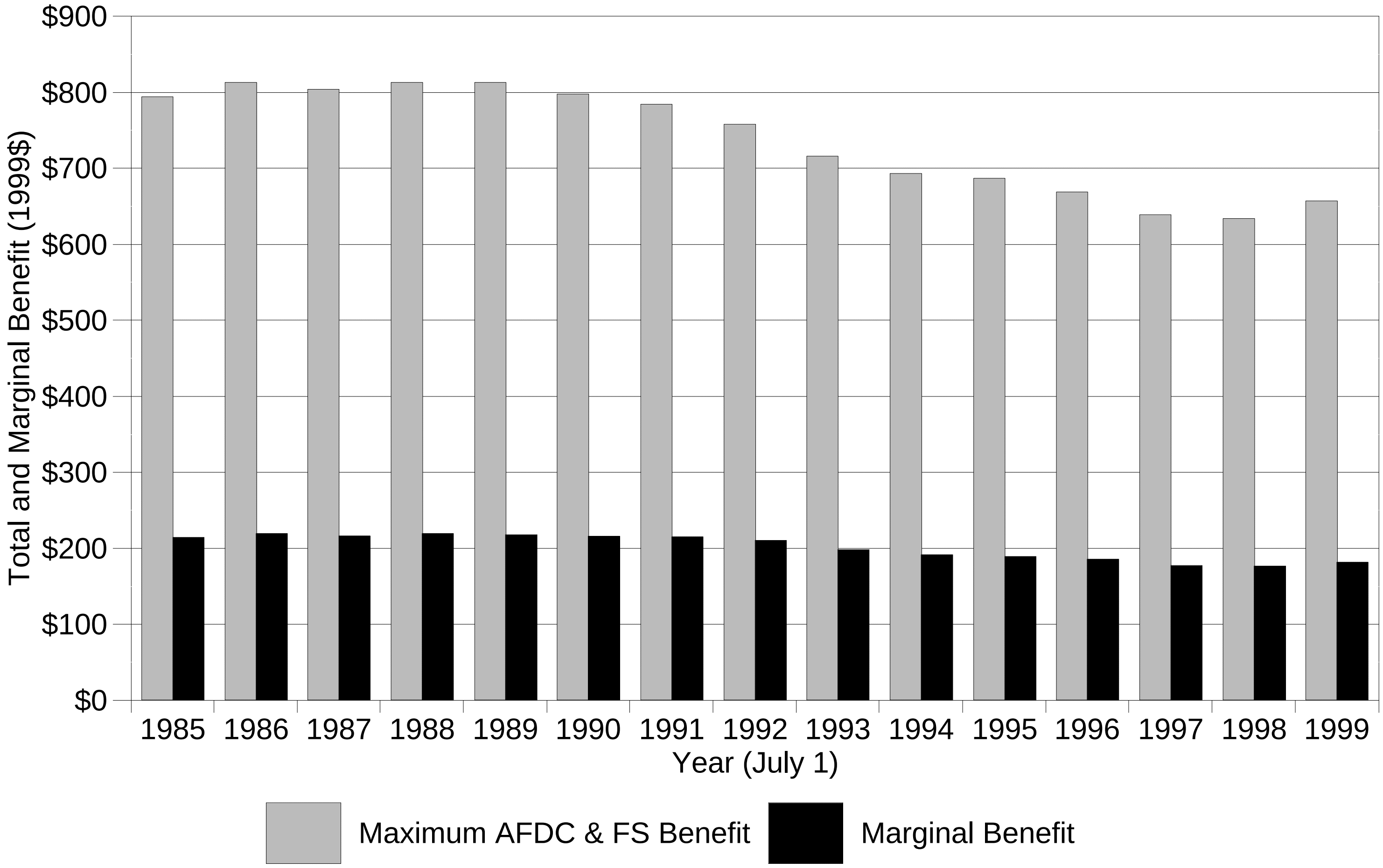
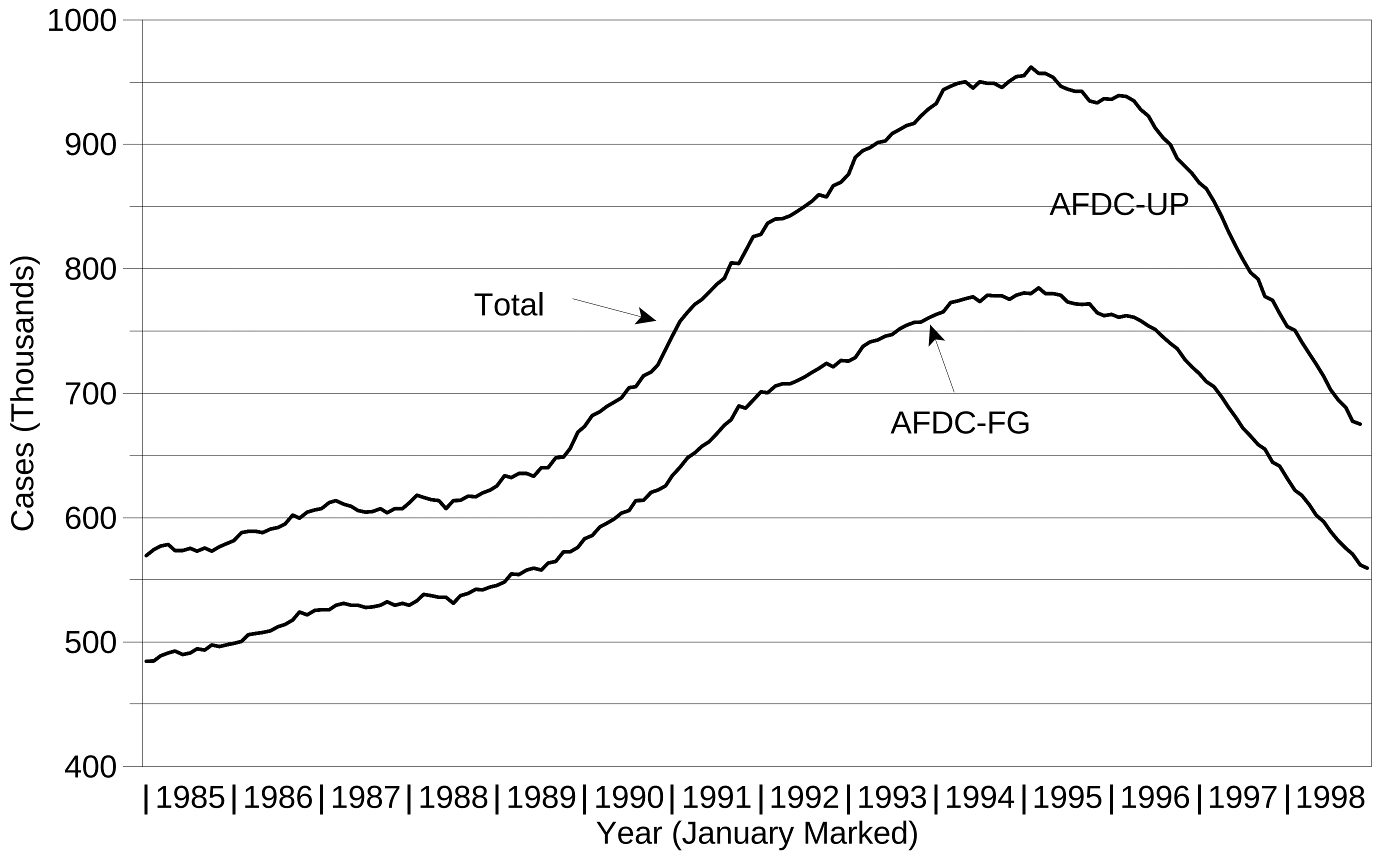


FIGURE 2
The California AFDC/TANF Caseload
1985B1998



in families receiving assistance from AFDC. The caseload expansion was associated with substantial demographic change. For example, 29 percent of AFDC-FG cases and 26 percent of AFDC-UP cases ever open during 1988 were classed as “Hispanic.” By 1995, these percentages had increased to 39 and 48, respectively (see Table 3, below). The state’s economic recovery from the 1991–92 recession lagged behind the rest of the country. Elsewhere the caseload turnaround began in late 1993 or early 1994 (the national peak was in March 1994). In California the turnaround began only in 1995, but subsequent reductions have been swift.

In summary, unchanged eligibility, falling benefits, and sharply rising caseloads form the context for studying the incidence of births among recipient families. Between 1988 and 1996, the number of children in AFDC families increased by over 600,000. How many of these children would have been subject to a benefit cap? The data used for answering this numbers question are introduced in the next section.

THE DATA

The numbers question seems simple enough in abstract, but it is difficult to answer in California. Unlike most other states, California does not have a centralized management information system for public assistance programs. Instead, the operating systems for AFDC/TANF eligibility and benefits vary across counties. Fortunately, over the past eight years the state has used an alternative source to develop a longitudinal public assistance database from which the number of births subject to the cap can be estimated. The basis is not the AFDC/TANF system itself, but rather the Medi-Cal Eligibility Data System (MEDS). MEDS data are reported by counties in conjunction with claiming reimbursement by the state for costs covered by Medi-Cal. When people are determined to be eligible for AFDC (and now TANF), they are automatically eligible for Medi-Cal. The data entered into the MEDS system include their AFDC case number. When people leave AFDC, their Medi-Cal status changes, even in the common situations in which Medi-Cal eligibility endures. These status codes allow separate identification of AFDC cases included among the many programs represented among Medi-Cal eligibles. As a result, MEDS is a source of information on comings and goings in all parts of California's public assistance program.

As is often the case when administrative data are used for social science, there are devils in the details. This section describes features of the MEDS data, the procedures used by the state to extract data from MEDS for policy studies, and the way in which these features have influenced the method of the present study. The conclusion is that both the incentives that influence data entry and the procedures for making the extract create problems, but that the data are nonetheless useful and offer a reliable platform for studying welfare conceptions. The principal caveat is that data from the last months of the analysis extract are most suspect, and caution is advised—and exercised in the rest of the paper—in dealing with information from this period.

The Longitudinal Database

The first step in the state's effort to move MEDS data from transactions to research support was creation of an extract. The Longitudinal Database of Cases (LDB) is a 10 percent sample of all cases reported as Medi-Cal eligible for at least one month in the interval 1987–1996 (UCDATA, 1997). The data cover the aid history of each person associated at any time with the sampled cases from January 1987, or the date of first receipt of public assistance anywhere in the state, whichever is later. The LDB was intended to be part of the state's data infrastructure for welfare policy research and evaluation of welfare reform initiatives. Public use copies of the LDB are made available to researchers by the state's contractor (the Survey Research Center at the University of California at Berkeley) subject to protocols designed to protect the confidentiality of the information contained.

The LDB consists of what are termed demographic and yearly data files. A demographic record exists for every person who appears in the LDB sample. The demographic file includes the case number, a person code unique to the individual, date of birth, ethnicity, sex, and certain information on case status at the point of first appearance of the person in the sample. A yearly data record exists for every person who appears in the LDB sample for every year between 1987 and 1996. Each year's observation includes the individual's person code and, for each month of the year, a code for type of assistance received and county of residence if aid is reported. Cases may be reconstructed from these person data by aggregating on the case number. The database evolves with the caseload as each annual cohort of new case openings is sampled. "New cases" in this instance means cases never open between January 1, 1987 and the date of the new cohort. The latest LDB case sample available is for 1996; with this cohort the database includes 1.4 million people and 625, 000 cases.

Problems

Given its size, statewide coverage, and longitudinal character, the MEDS/LDB would appear ideal as a resource for evaluating the consequences of the cap. However, closer inspection reveals shortcomings with regard both to identification of parents and data accuracy.

The problem with identifying parents arises because of the way eligibility for Medi-Cal is recorded. Eligibility for Medi-Cal is determined on an individual, rather than a family or household, basis. The MEDS file includes information on the various programs for which each person in a case is eligible, but it does not include information on family structure. Users cannot tell, for example, if the oldest woman in a case is or is not the mother of the youngest child. Often the mother's and the father's identities can be reasonably inferred, as in the case of an AFDC-UP family that includes only one adult female and one adult male both aged less than 30 and only preschool children. But some cases include both older and younger women, and in such instances if a baby appears it is impossible to determine from the MEDS file to whom the child was born. In California a substantial number of adults are not eligible for AFDC assistance because they are not citizens. However, if they are needy and if the children were born in the country, they are eligible for benefits on the children's behalf. Since in such cases the mother is ineligible, the MEDS file includes nothing on her characteristics.

Beyond shortcomings arising from the conventions of Medi-Cal eligibility, problems are generated by the county agencies responsible for data upkeep. The codebook distributed with the LDB data emphasizes that the MEDS records are updated through a statewide electronic network and that changes occur daily (UCDATA, 1997, p. 25). The implication is that changes in program status are quickly and accurately reported. In fact, nuances of AFDC status are of little relevance to county reimbursement for Medi-Cal costs; the Medicaid program makes no distinction between, to take an important example, AFDC-FG and AFDC-UP. From a county's perspective, what counts is to get Medi-Cal eligibles into a reimbursable category. There are no penalties for mislabeling, just for claiming reimbursement for expenses incurred for persons or families not eligible for assistance under any category. Thus identification errors are likely to occur.

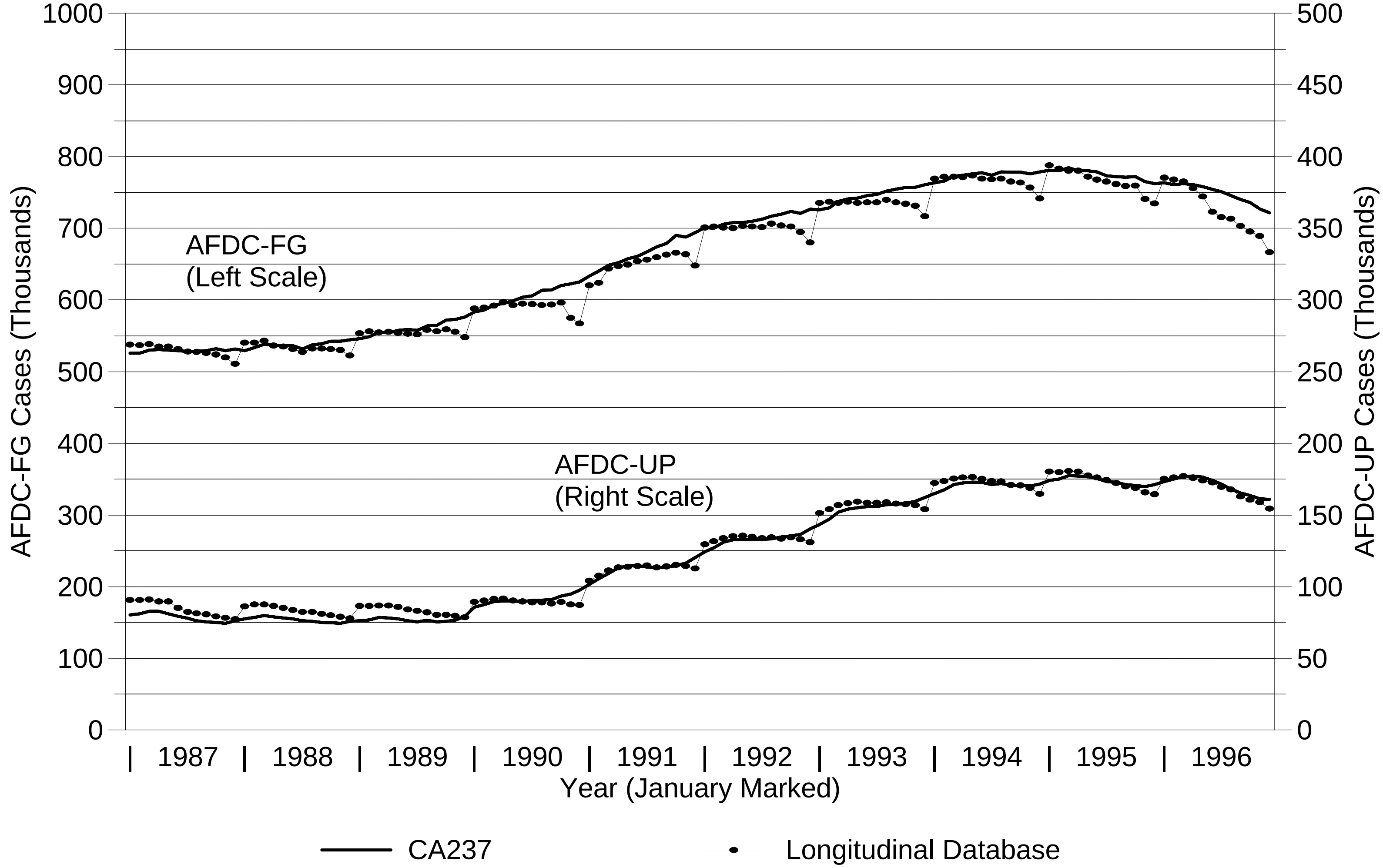
In addition to possible problems in program assignment, the MEDS and LDB extract data can deviate from official timing of entry and exit because of problems associated with data entry into the MEDS system. For most counties over most of the history covered by the system, MEDS entry has involved a separate system and often separate terminals from the program used for AFDC eligibility. As a result, a case opened in one month may not get into the MEDS system until one or two months later, and adjustments after closure may also occur with a lag. When adjustment does occur, the newly entered data should be the actual dates of entry, exit, or program change, but every adjustment provides opportunity for error. In practice, the incentives for getting new cases into the MEDS file are stronger than the incentives for prompt adjustment of cases on closure. Counties have substantial fiscal involvement with the hospital and physicians who provide Medi-Cal services, and reimbursement for the state and federal share of costs can only be accomplished for persons registered as eligible in the system. Slowdown in both data update and entry are particularly evident in the last quarter of the year.

Lags in MEDS data entry and update interact with LDB extract to reduce data quality for the last months of the most recent year of data. MEDS data for the LDB are extracted from an annual extract archived in January for the previous January–December transactions records. Data for the early months of the year have a much greater chance of being updated and corrected—indeed even entered—than data for the last months (UCDATA, 1997, p. 7). The result is that the last months of the latest available LDB extract tend to be the least reliable. Subsequent data extracts lead to updates of earlier intervals, so, for example, data for 1995 changed with the January 1997 extract of information for 1996.

This result is apparent in Figure 3, where caseload counts for AFDC-FG and AFDC-UP derived from the LDB are compared with the AFDC caseloads reported monthly by counties directly to the state. These administrative caseload reports, called CA237 after the name of the form on which they are submitted, are the basis for all state reporting to the federal government and the state's own published caseload summaries. The CA237 data are the most reliable for caseload, since they come straight from county payment systems. The good news is that the LDB does track the CA237 caseload over the long span of the sample. The bad news is that local agencies lag in updating their MEDS files near the end of the year. For AFDC-FG it is clear that all updating had not been accomplished by the time of the last (1997) extract; if past experience is replicated, the data for the 1996 extract (made in January 1997) would have been modified had an extract of 1997 MEDS records been made in 1998.

The concern of this paper is the fertility behavior of adults already receiving welfare, and therefore failure to quickly record accessions or changes from one component of the AFDC program to the other, on the scale evident in Figure 3, is unlikely to be a problem. However, the focus is on a change—a birth. If the MEDS file fails to capture infants added to families receiving AFDC, it is of no use for studying the likely effect of benefit caps. To assess the responsiveness of MEDS reporting to births, a tabulation was conducted of all children appearing in the MEDS file before their first birthday who were born into families receiving AFDC at the time and whose first reported aid code was AFDC-

FIGURE 3
Caseload Comparison, LDB versus CA237
1988B1996



FG or AFDC-UP. Virtually all of these babies made it into the MEDS file within six months of birth. In addition, no change was evident in the distribution of time between birth and MEDS/LDB addition over the years of the sample. It is reasonable to assume that such infants are in fact included in the AFDC family budget unit from birth, and that it is addition to the MEDS file that is slow to occur. In the analysis that follows, any infant is classed as "born on public assistance" if it is added to the MEDS file as an AFDC recipient within six months of birth and if others in the same case were receiving AFDC at the actual date of birth. This restriction has an important consequence for the span of the study. Because of the lag in MEDS update, presumably some of the infants born in 1996 do not show up in the LDB until 1997 (and no 1997 case sample is available). Therefore the last year included in the study of births is 1995, since this is the last year available for which data have gone through both an original extract and an update. Note that infants born on assistance include cases receiving money only on behalf of children, and there is no specific link between children who meet these criteria and adults who were in the associated case at the probable time of conception. Also, the data do not permit adjustment for premature births.

In summary, the LDB has many attractive features for studying the likely effect of a benefit cap, but as in other applications of administrative data for social science, it is important to understand the process by which the data are generated and the motivation of those who assemble it. Special caution must be exercised in analyzing the data from the last of the sequence of extracts.

COUNTING WELFARE'S CHILDREN

Results

The results for 1988 and 1995 are summarized in Table 2. The table is divided between information on case and birth counts and information about family composition, and separate results are presented for AFDC-FG and AFDC-UP.

TABLE 2
Births to Welfare Recipient Families
California, 1988 and 1995

		AFDC-FG		AFDC-UP	
		1988	1995	1988	1995
Case Counts					
(a)	Average beginning-of-month caseload	504,906	736,189	70,960	163,772
(b)	Cases ever open during calendar year	701,710	959,710	130,240	244,320
	Turnover ratio [(b)/(a); see text]	1.39	1.30	1.84	1.49
(c)	Cases with births	68,420	74,840	16,020	23,380
(d)	Ratio (c)/(b), the incidence of cases with births among all cases open during the year	0.098	0.078	0.123	0.096
(e)	Cases with newborns conceived on AFDC	36,280	45,680	10,400	16,420
(f)	Ratio (e)/(c), the share of births attributable to infants conceived on welfare	0.530	0.610	0.649	0.702
(g)	Product (f)*(d), the incidence of births of children conceived on welfare among all cases ever open during year	0.052	0.048	0.080	0.067
Case Composition at Time of Birth					
	Proportion of cases:				
(h)	with female adult in assistance unit, no teen female present	0.831	0.781	0.883	0.737
(i)	with female adult in assistance unit, teen female (aged 15–17) also present	0.035	0.041	0.034	0.044
(j)	without female adult in assistance unit, no teen female present	0.075	0.117	0.036	0.183
(k)	without female adult in assistance unit, teen female present	0.053	0.050	0.037	0.030
(l)	with two adult females on budget	0.006	0.011	0.009	0.005

Source: Calculations by author from California Longitudinal Database [UCDATA, 1997].

The first number in each column is the average beginning-of-month caseload for the subprogram and year indicated. The second row reports the estimated total number of cases ever receiving benefit within the subprogram over the entire year. (The number is simply ten times the case count from the LDB; note that each number in the second row ends with zero.) One measure of turnover is the ratio of cases ever open during a year to the average monthly caseload. This is the ratio of the program number in line (b) to the number in line (a). As is to be expected, turnover in the AFDC-UP program (ratio 1.84 in 1988) is greater than in AFDC-FG (ratio 1.39); turnover in both programs was lower in 1995 than in 1988. The lower turnover in 1995 is consistent with the argument of the previous section that the caseload overestimate within the LDB is attributable to turnover.

Line (c) reports the number of cases with births (a multiple birth counts as one case with birth), and line (d) is the incidence of birthing among cases. In what follows, the term *rate* will be used intermittently with *incidence*. Formally speaking, this convention is not quite appropriate. As used in demography, the birth rate would cover all births to these families within the calendar year regardless of assistance status, while the LDB provides data only on births charged to Medi-Cal and the table counts only those that occurred while the family's AFDC case was open. Overall in 1988 about one in ten cases open at any time of the year is associated with a newborn. This rate is higher for AFDC-UP cases (.l23) than it is for AFDC-FG (.098). All rates appear to have fallen between 1988 and 1995. Changes of this order are statistically significant, given the large sample sizes.

Line (e) is the subset of the births (c) that satisfy the welfare conception criteria. Both the absolute number of such births and the share of such births among all births to families receiving assistance increased between 1988 and 1995, as seen in line (f). Statewide, in 1996 some 48,000 cases included births that would have been subject to a benefit cap had the state implemented such a policy. While the share of total births attributable to welfare conceptions has risen, the incidence of such births across cases has fallen. Line (g) shows the change, and the result is a major discovery of this analysis:

Between 1988 and 1995 the incidence of births among welfare recipients in California that were the result of conceptions occurring while the mother was receiving assistance fell by 8 percent in AFDC-FG and 16 percent in AFDC-UP.

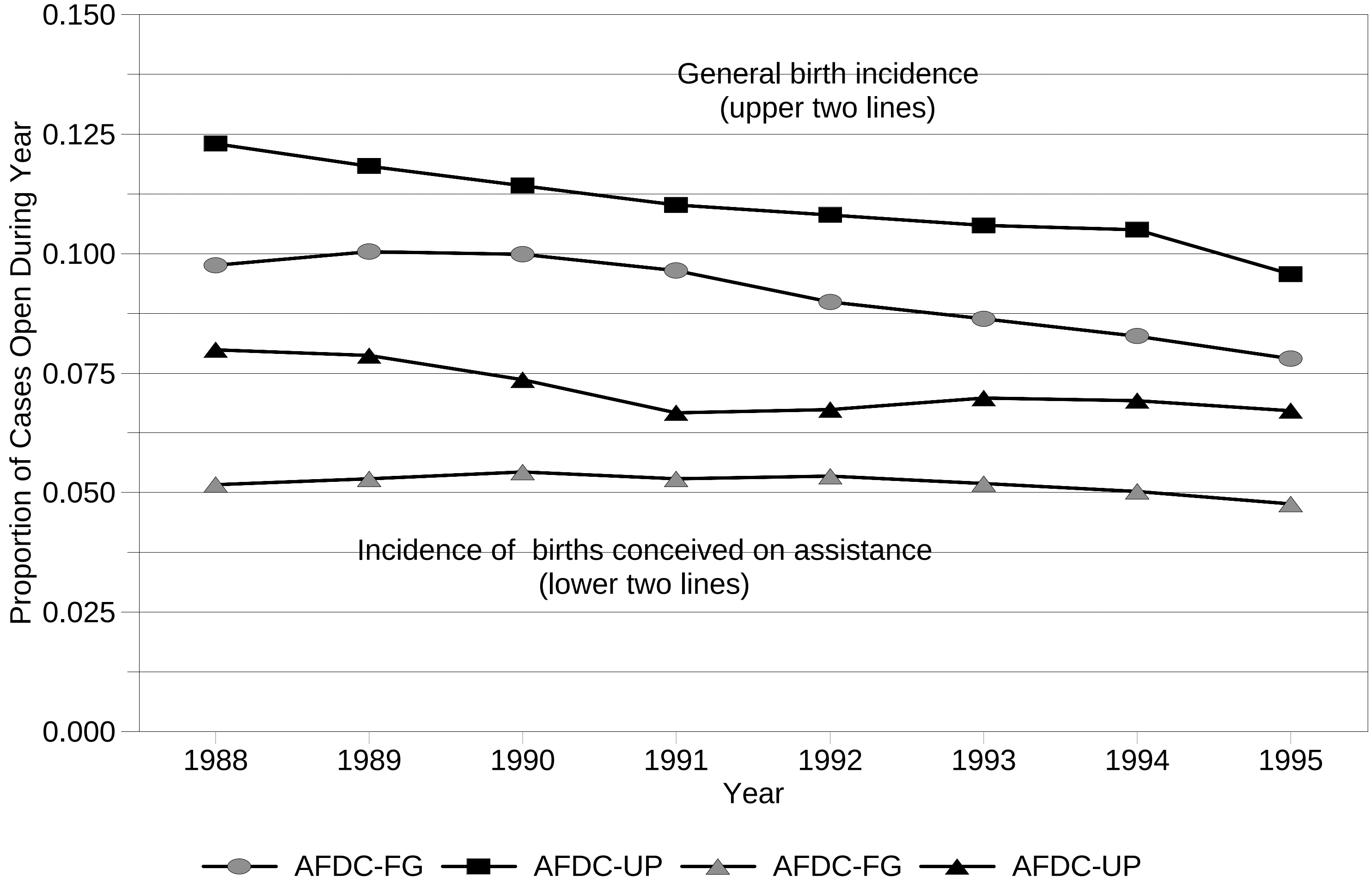
As discussed earlier, given the nature of the MEDS data, there is some ambiguity concerning parenting. Lines (h)–(l) provide information on the composition of cases with births. In 1988, 83 percent of AFDC-FG and 88 percent of AFDC-UP cases were unambiguous in the sense that the cases included an adult female and there was no teen female aged 15–17 reported (the small number of births to teens younger than 15 is ignored here). Both of these proportions declined between 1988 and 1996, principally because the share of cases without an adult female or teenage recipient aged 15–17 grew. Parents are definitely present in the households containing these cases (otherwise the case would be classified differently), but they are ineligible for AFDC benefits either because they are not citizens or because they are receiving Supplemental Security Income. Regardless of the parents' own welfare status, today such cases would be subject to California's family cap.

The bottom section of Table 2 shows case composition at time of birth. In about 9 percent of AFDC-FG cases and 7–8 percent of AFDC-UP cases, the parenting of new infants is uncertain because of the presence of a teenage female less than 18 years old on a case. However, it is not the incidence of potential mothers among children in some cases that has changed but rather the frequency of absence of an adult. The share of cases without adults counted in computing the budget went up substantially during this period in both AFDC-FG and AFDC-UP.

<u>Trends</u>

Figure 4 reports the trend in the ratios reported in lines (d) and (g) of Table 2. The birth incidence data in the upper portion of the chart indicate that birth rates began to fall in 1991–92. The trends coincide with trends in birth rates nationwide (Ventura et al., 2000). Nationwide, for women aged 15–29 with one child, the incidence of second births (roughly 190 per 1,000) peaked in 1990–91. Both the birth rate for all women 15–44 and the rate for women 15–19 peaked in 1990. Also nationally, the

FIGURE 4
Birth Incidence, Recipient Families
by Conception Status, 1988B1995



rate of nonmarital births peaked (at 45.9 per 1,000 unmarried women) in 1994. Although the birth incidence among generally unmarried AFDC-FG mothers receiving assistance in California is higher (83 per 1,000 in that same year), the decline started earlier.

Though the overall incidence of births among assistance cases is clearly declining, the results for births of children conceived on assistance are ambiguous. In AFDC-FG, rates rose slightly after 1988, then began a gradual decline in 1992. In AFDC-UP, this rate rose slightly between 1991 and 1995. The figure indicates that, as was the case for AFDC-FG, the 1988–1995 change presented in Table 2 is not the product of a consistent downward trend.

For AFDC-UP there may in fact be no downward trend at all, just a change in mix. Birth rates and the incidence of welfare conceptions differ across race and ethnic groups. The LDB includes an extensive race/ethnicity breakdown (UCDATA, 1997, p. 17). Some groups—Filipinos, for example—have too few recipients for separate analysis. This paper focuses on six groups. Using the same descriptors applied by the California Department of Social Services, these are (in reverse order of size), white (non-Hispanic), Hispanic, black, Vietnamese, Laotian, and Cambodian. All other groups, including the relatively few cases for which race/ethnicity was not identified, are combined as Other.

Table 3 reports the equivalent of the data in lines (e) and (g) from Table 2 for each of the race/ethnicity groups. While levels differ substantially, for AFDC-FG cases the trends within each group match that apparent in the aggregate data and the 1988–1995 change is statistically significant for Hispanic, black, Cambodian, and Vietnamese families. For AFDC-UP the results provide a good example of the perils of aggregation. For *every* group except the Vietnamese (7 percent of the 1995 caseload) and Other (4 percent of the 1995 caseload) category families, the incidence of births of children conceived on welfare went up, but the aggregate incidence went down. The reason is evident from the numbers at the top of the table. Between 1988 and 1995 the share of the AFDC-UP caseload attributable to Hispanic

TABLE 3
Variation in Births and Potential Cap Effects by Race/Ethnicity
California, 1988 and 1995

	AFDC-FG				AFDC-UP			
	1988	1995	Change	P(z)**	1988	1995	Change	P(z)**
Cases ever open during year	701,710	959,710	258,000		130,240	244,320	114,080	
Share of cases ever open*								
Total	1.000	1.000			1.000	1.000	0.000	
White (non-Hispanic)	0.385	0.317	-0.068		0.384	0.304	-0.079	
Hispanic	0.291	0.393	0.103		0.258	0.475	0.217	
Black	0.252	0.210	-0.042		0.065	0.047	-0.018	
Cambodian	0.005	0.011	0.006		0.023	0.010	-0.012	
Laotian	0.005	0.011	0.006		0.056	0.031	-0.024	
Vietnamese	0.018	0.020	0.002		0.088	0.072	-0.016	
Other	0.034	0.026	-0.008		0.097	0.040	-0.057	
Incidence of births in cases ever open								
Total	0.098	0.078	-0.020	0.00	0.123	0.096	-0.027	0.00
White (non-Hispanic)	0.082	0.067	-0.015	0.00	0.114	0.095	-0.019	0.00
Hispanic	0.100	0.080	-0.020	0.00	0.092	0.089	-0.003	0.63
Black	0.120	0.093	-0.027	0.00	0.126	0.114	-0.012	0.41
Cambodian	0.229	0.078	-0.152	0.00	0.125	0.131	0.006	0.85
Laotian	0.143	0.089	-0.054	0.00	0.247	0.243	-0.004	0.87
Vietnamese	0.066	0.059	-0.007	0.40	0.128	0.073	-0.055	0.00
Other	0.087	0.080	-0.006	0.42	0.178	0.093	-0.085	0.00

(table continues)

TABLE 3, continued

	AFDC-FG				AFDC-UP			
	1988	1995	Change	P(z)**	1988	1995	Change	P(z)**
Incidence of births potentially subject to cap								
Total	0.052	0.048	-0.004	0.00	0.080	0.067	-0.013	0.00
White (non-Hispanic)	0.035	0.034	-0.001	0.69	0.058	0.060	0.002	0.69
Hispanic	0.059	0.053	-0.006	0.00	0.054	0.061	0.007	0.13
Black	0.070	0.060	-0.011	0.00	0.079	0.090	0.011	0.39
Cambodian	0.125	0.060	-0.065	0.00	0.112	0.119	0.007	0.79
Laotian	0.097	0.074	-0.023	0.16	0.209	0.213	0.005	0.82
Vietnamese	0.045	0.031	-0.013	0.05	0.104	0.062	-0.042	0.00
Other	0.045	0.044	-0.001	0.89	0.142	0.065	-0.077	0.00

Source: Calculations by author from California Longitudinal Database [UCDATA, 1997].

*Proportions do not sum to 1 because of a small number of cases with missing race/ethnicity data.

**P(z) is the chance that a difference between birth incidence in 1995 and 1988 as large or larger than calculated here using the LDB case sample might result from sampling variability given that the true incidence was the same in both years.

families nearly doubled, while the shares of all other groups declined. Welfare conceptions are relatively uncommon in Hispanic AFDC-UP cases.

Table 3 indicates that the demographic composition of the caseload changed substantially over the eight years from 1988 to 1996. Other data, not shown here, reveal that change in racial/ethnic composition is not the only development in the demography of the AFDC caseload over this period. Overall the average age of mothers has increased, and changes have occurred as well in the size of cases and time on welfare. This suggests that the message of Table 2 is possibly confused by the interaction of mixture effects and behavioral change. The next section adds more control through multivariate analysis, with interesting effect.

MULTIVARIATE ANALYSIS OF FERTILITY

To this point the discussion has been framed in aggregates, dealing with matters such as all families receiving benefit, all children born during the year, and so forth. This section moves beyond the incidence of births subject to the cap to study family characteristics associated with the conceptions that lead to such births and changes over time in the incidence of such conceptions holding family characteristics constant.

The sample to be analyzed consists of all open cases in February (called here the "reference month") of each year 1988–1995 in which there is no pregnant woman who will bear a child over the coming nine months. This, therefore, is a point-in-time sample, unlike the data used for Table 2 and Table 3, which covered all cases ever opened over a time interval—a year. The target is the likelihood that over the course of the next 12 months some female in the case will conceive a child that will be carried to term, born, and added to the family assistance unit—in other words, a child satisfying the criteria for "conceived on AFDC" in Table 2. What factors are associated with this outcome? This question is approached by estimating the coefficients of a multivariate probit model. The dependent variable is the inverse of the cumulative standard normal distribution evaluated at the probability of a conception that will lead to application of the benefit cap. Separate equations are estimated for AFDC-FG and AFDC-UP (in this case, program is defined by case status as of February, not as of the child's birth).[2] The right-hand variables include measures of the age of the oldest adult female in the case (plus an indicator for cases in which there is no adult female receiving assistance), the number of children, children's ages, and ethnicity. Because of left-censoring it is not possible to control for total time on assistance, but the model does include two indicators for case status over the 12 months preceding the reference month. The independent variables are defined in the appendix.

Results

AFDC-FG. The estimated coefficients for both AFDC-FG and AFDC-UP are reported in Table 4. Considering first the equation for AFDC-FG and examining only the relative size of coefficients and statistical significance of the outcomes, the following conclusions, all stated ceteris paribus, seem statistically justified and important:

- **In cases that include teenage or adult women, the greatest chance of conception occurs within the age groups 15–17 and 18–19.**

Recall that we cannot be certain that those cases with teenagers 15–17 do not include an older female who is not part of the family budget unit. All age variables for cases without an identified female 15 or older are zero; for this group NOFEMALE = 1.

- **African-American and Hispanic cases are significantly more likely to experience welfare conceptions than are white ones; Cambodian and Laotian cases are significantly more likely to record such conceptions than are Hispanics, blacks, or whites.**

- **Women in AFDC-FG who have had two children are more likely to conceive than are women in cases which include only one child.**

[2]The reported models include no adjustment for county fixed effects or the presence of multiple observations on households appearing in the data for more than one year. Estimation with county fixed effects and making the Huber-White variance adjustment does not significantly affect either the coefficient estimates themselves or their estimated standard errors. These estimates are available from the author.

TABLE 4
Probit Models, Probability of Welfare Conception

Variable[a]	AFDC-FG (Sample Mean Rate = .061) Sample Mean	Coefficient	Standard Error	P>|z|[b]	AFDC-UP (Sample Mean Rate = 0.079) Sample Mean	Coefficient	Standard Error	P>|z|[b]
AGE15-17	0.036	0.6440	0.0168	0.000	0.020	0.6240	0.0449	0.000
AGE18-19	0.035	0.6702	0.0169	0.000	0.022	0.6326	0.0418	0.000
AGE20-24	0.148	0.5198	0.0127	0.000	0.117	0.5701	0.0264	0.000
AGE25-29	0.163	0.3692	0.0123	0.000	0.162	0.4037	0.0240	0.000
AGE30-34	0.149	0.2288	0.0127	0.000	0.174	0.2608	0.0238	0.000
NOFEMALE	0.268	0.1358	0.0121	0.000	0.202	0.2329	0.0264	0.000
ETHBLACK	0.241	0.2224	0.0082	0.000	0.044	0.1089	0.0324	0.001
ETHCAMB	0.010	0.3487	0.0278	0.000	0.017	0.4153	0.0462	0.000
ETHHISP	0.350	0.2524	0.0078	0.000	0.366	0.0952	0.0174	0.000
ETHLAOT	0.010	0.4169	0.0277	0.000	0.050	0.6559	0.0265	0.000
ETHVIET	0.022	0.1156	0.0234	0.000	0.110	0.2427	0.0249	0.000
ETHOTHER	0.007	-0.0555	0.0456	0.224	0.014	-0.0950	0.0659	0.150
CHILD>2	0.526	0.0166	0.0074	0.025	0.745	-0.1020	0.0188	0.000
CHILD>3	0.224	-0.0110	0.0099	0.263	0.434	-0.0569	0.0197	0.004
CHILD>4	0.820	0.1005	0.0128	0.000	0.218	0.1376	0.0208	0.000
YNGSTCHL<2	0.275	0.0830	0.0097	0.000	0.374	0.1790	0.0207	0.000
YNGSTCHL<3	0.381	0.1327	0.0105	0.000	0.492	0.1563	0.0240	0.000
YNGSTCHL<6	0.608	0.2604	0.0096	0.000	0.703	0.2995	0.0249	0.000
YNGSTCHL>15	0.033	0.0341	0.0261	0.192	0.031	-0.2924	0.0795	0.000
ONFEB	0.060	0.0440	0.0133	0.001	0.055	0.0326	0.0303	0.281
ONALL	0.631	0.0581	0.0069	0.000	0.575	0.0939	0.0156	0.000

(table continues)

TABLE 4, continued

Variable[a]	AFDC-FG (Sample Mean Rate = .061) Sample Mean	Coefficient	Standard Error	P>|z|[b]	AFDC-UP (Sample Mean Rate = 0.079) Sample Mean	Coefficient	Standard Error	P>|z|[b]
YEAR>1989	0.898	-0.0130	0.0134	0.331	0.916	-0.1225	0.0319	0.000
YEAR>1990	0.794	0.0502	0.0130	0.000	0.832	0.0799	0.0320	0.012
YEAR>1991	0.685	-0.0029	0.0124	0.818	0.744	-0.0056	0.0297	0.849
YEAR>1992	0.568	-0.0143	0.0119	0.229	0.639	-0.0597	0.0273	0.028
YEAR>1993	0.436	-0.0291	0.0115	0.012	0.509	0.0230	0.0251	0.358
YEAR>1994	0.296	-0.0297	0.0115	0.010	0.356	-0.0380	0.0235	0.106
YEAR>1995	0.150	-0.0608	0.0116	0.000	0.182	-0.0893	0.0234	0.000
INTERCEPT	1.000	-2.3271	0.0158	0.000	1.000	-2.1264	0.0349	0.000
Number of observations		461,014				83,429		
Log likelihood		-98894				-20943		
Chi-square (28)		13,579				4,112		
Prob > chi-square		0.000				0.000		
Pseudo R^2		0.064				0.089		

[a]For variable definitions, see the appendix.

[b]z is the ratio of the coefficient estimate to its standard error. P is the probability of observing a ratio with absolute value this high or higher in a sample of this size drawn from a universe in which the true coefficient is zero.

Very large case size, that is cases with four or more children, is associated with greater likelihood of an additional conception. Thus to some extent these probits confirm a stereotype: large families are associated with dependency-enhancing childbearing.

- **The younger the youngest child in the household, the greater the risk of another conception.**

The first three "young child" variables are coded sequentially, so that, for example, if the youngest child in the case is less than 2 years old, all three variables YNGSTCHL<2, YNGSTCHL<3, and YNGSTCHL<6 will be coded "1," and every estimated coefficient is positive and statistically significant. On the other hand, AFDC-FG cases at risk of loss of eligibility due to aging out of the youngest child (YNGSTCHL≥15 = 1) are not significantly more likely to report a conception than are others.

- **Women who have been receiving assistance for less than 12 months at the reference month are less likely to conceive over the coming year than are others.**

ONFEB and ONALL are mutually exclusive; the difference in coefficient size is not statistically significant.

So far, the multivariate analysis for AFDC-FG cases has yielded few surprises. If there is something dramatic here, it lies in the time profile of the conception probability. The probability of conception peaks in 1991 and then begins to fall.

> **Case data drawn from the California Longitudinal Database indicate that the likelihood that a mother in an AFDC-FG case will conceive an additional child while receiving benefits reached a maximum in 1990 and then declined through the next four years.**

This decline substantially predates the implementation of California's benefit cap, but not the public discussion of such policies. Note that given the way the observations are coded, the coefficient for each year indicator is the *change* in predicted outcome given the move from the preceding year's subsample to that registered by the preceding variable. The associated *t*-test is therefore a test of the null hypothesis that, in moving from the preceding to the current year, no change occurred in the likelihood of welfare conception.

AFDC-UP. The results for AFDC-UP are reported on the right side of Table 4. As expected, the behavior of two-parent families is in some ways different from that of single parents. Again, the age group at greatest risk is 15–19, although the difference in coefficients for all three age groups 15–17, 18–19, and 20–24 is not statistically significant. Here as with AFDC-FG, African-American and Hispanic cases are significantly more likely to experience welfare conceptions than are white ones; Cambodian and Laotian cases are significantly more likely to record such conceptions than are Hispanics, blacks, or whites. Vietnamese cases fall in between. Here, too, the presence of very young children raises the likelihood of conception, but unlike the case for AFDC-FG, an increase in the number of children already born from one to two or from two to three is associated with a reduction, not an increase, in the likelihood of an additional conception. AFDC-UP couples who have received benefits for less than a year are less likely to conceive an additional child during the observation period than are others. The pattern of change over time is more ambiguous for AFDC-UP than for AFDC-FG. The coefficients for both 1989 and 1994 are negative and virtually identical. However, some decline is evident, and the rate for 1995 is significantly different from both the rate for 1994 and the rate for the initial observations in 1988.

Caution is appropriate in considering the 1995 coefficient in both the AFDC-FG and AFDC-UP models. Recall that the 1995 data were originally extracted from MEDS in January 1996 and were updated once in January 1997. Given the one-year horizon of the analysis of conception probability, it is possible that a birth might occur as late as September 1996. The September 1996 data are obtained only from the 1997 extract, and there is a lag in recording some births. Hence it is possible that the decline in welfare conceptions between 1994 and 1995 is exaggerated by an undercount. A test of this possibility was conducted by redefining the outcome of interest from a conception over the year subsequent to the February point of observation to conception over the subsequent six months; this would place all births in the interval November–April. As suspected, the proportionate reduction between 1994 and 1995 is reduced, but the trend is not eliminated. Given this outcome, the observations for 1995 are retained in the discussion that follows, but the downward bias in the estimated coefficient for this period should not be forgotten. The undercount issue is discussed again in the conclusions.

Probabilities

Probit coefficients are not expressed in a natural metric; the matter of policy interest is the effects of changes in the various independent variables on the probability of a welfare conception. Table 5 presents results of calculation of probability of conception for a sample woman aged 22, black, with one child aged 3, on welfare for more than a year, and observed in February 1991. (Note that this is not the reference group for the probit regression itself.) Each of these characteristics is then altered, and the original and revised probabilities are shown for each program. The calculations give dimension to the inferences drawn from the probits themselves. Note the following:

- **The probability of a welfare conception is higher for the two-parent families in AFDC-UP and substantially greater in cases with a mother younger than 20 than for cases with a mother over 25.**

The fact that AFDC-FG probabilities are lower than those for AFDC-UP does not seem surprising; the small size of the estimated difference may be.

- **The younger the first child, the more likely an additional conception becomes.**

If the first child in a case reaches age 3 without a younger sibling, the probability of a welfare conception in the following year falls by 30 percent (from .155 to .109) in AFDC-FG and by 40 percent (from .223 to .134) in AFDC-UP.

- **The estimated decline in the probability of conception over the first half of the decade is substantial, amounting to 21 percent in AFDC-FG (from .109 to .086) and 26 percent (from .136 to .101) in AFDC-UP.**

Clearly there is some uncertainty about the amount, since the results are not as dramatic if 1994 is compared to 1990 to avoid the question raised earlier about the 1995 coefficient. Nevertheless, virtually

TABLE 5
Estimated Probabilities for Welfare Conception

			AFDC-FG		AFDC-UP	
Characteristic	Reference Assumption	Change to	Revised Probability	Probability Change	Revised Probability	Probability Change
Age	20–24 years old		0.109	—	0.134	—
		18–19	0.139	0.031	0.149	0.014
		25–29	0.083	-0.025	0.103	-0.031
Race	Black		0.109	—	0.134	—
		White	0.072	-0.036	0.112	-0.023
		Hispanic	0.114	0.006	0.131	-0.003
		Cambodian	0.136	0.027	0.213	0.078
Family composition	One child, 3 years old		0.109	—	0.134	—
		Child age 2	0.136	0.027	0.172	0.038
		Child age 1	0.155	0.046	0.223	0.088
		Two children, additional child age 5	0.139	0.030	0.147	0.012
		1988	0.102	-0.006	0.145	0.010
		1989	0.100	-0.009	0.119	-0.016
		1990	0.109	0.001	0.136	0.001
Year	1991	1991	0.109	—	0.134	—
		1992	0.106	-0.003	0.122	-0.013
		1993	0.101	-0.008	0.126	-0.008
		1994	0.096	-0.013	0.118	-0.016
		1995	0.086	-0.023	0.101	-0.033

Source: Calculated from probit equations reported in Table 4.

*The reference case for these calculations is an African-American family residing in Alameda County. The case has been open continuously for a year or more.

any cut shows some decline. Thus the changes evident in the aggregate interval data hold up when control for demography is added.

Racial Variation

As a last check, the basic probit model was reestimated separately for each of the six racial/ethnic groups for both programs. Figure 5 plots the estimated probability of conception for a representative woman in each group by year. These estimates were constructed in exactly the same way and under the same assumptions as were the representative case estimates reported in Table 5, except that the estimated time trend is based on a separate probit regression for each group. The plot shows the probability by year. For black, Hispanic, Cambodian, and Vietnamese families, separate estimation confirms the downward trend since 1990–91 evident in the aggregate regression, and the results for Laotians are stronger. The incidence of welfare conceptions among white families shows little change over the entire interval.

For AFDC-UP, things are, again, more ambiguous. Figure 6 reports results comparable to those in Figure 5 for two-parent households. There is some downward trend for each group, but trend perception is very much influenced by the data for 1995. Recall the concern expressed about the 1995 data. Without the 1995 effect, there is little evident downward trend for Hispanic, black, or white families.

WHAT WOULD A BENEFIT CAP COST? HOW MUCH WOULD IT SAVE?

In fiscal year 1996, AFDC in California cost $6.5 billion in benefits and administrative costs (Committee on Ways and Means, 1998, pp. 422, 432). Suppose that AFDC had always included a cap. How much less would this bill be?

FIGURE 5
AFDC-FG Conception Probability
by Race/Ethnicity, California, 1988B1995

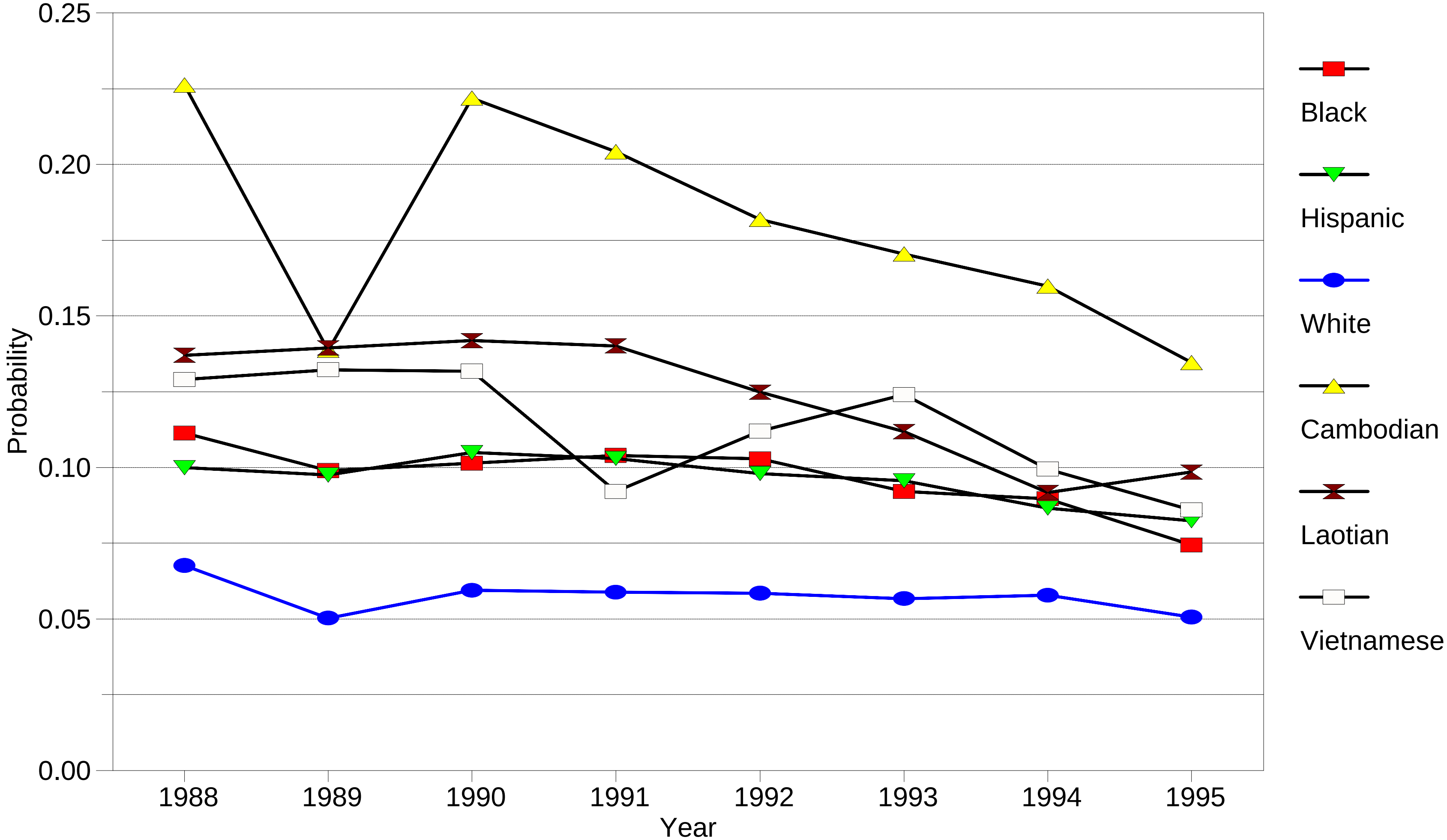
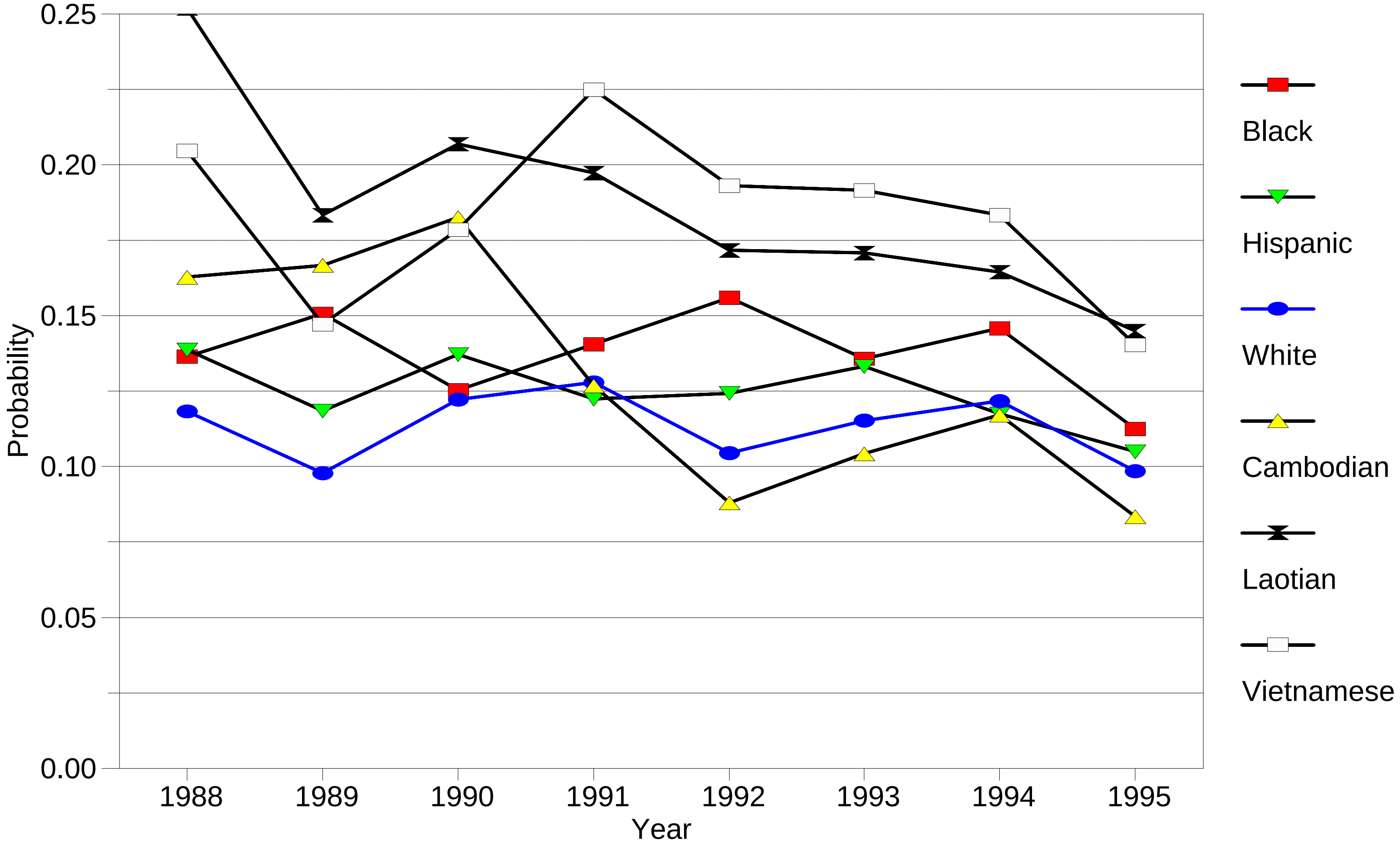


FIGURE 6
AFDC-UP Conception Probability
by Race/Ethnicity, California, 1988B1995



It is not possible to say with precision, for at least three reasons. One is that the LDB does not extend back far enough in time to establish the status at conception of the parents of each child on assistance. The LDB begins in January 1987, and this means that, given a ten-month grace period the applicability of a cap can be judged only for children born after November 1987. In the last year of the data, these children are at most 9 years old. A second reason is that the literature does not provide enough information to incorporate the behavioral effects of a cap once instituted. Most models of caseload dynamics suggest that, other things equal, smaller cases exit faster than large ones. If a cap reduces fertility, it would therefore also reduce the expected duration of cases. The public would save both through smaller benefit payments and because fewer cases would be open at any time.

The third reason has to do with details of the cap actually applied in California. The cap, termed the Maximum Family Grant (MFG) Rule, was established by law in 1994 on the basis of a previously granted federal waiver (California Department of Social Services, 1996). Application began August 1, 1997; notices were sent to recipient families in July 1996 that cautioned that the MFG rule could apply to any child born after July 1, 1997. The legislation allowed for aid to be paid on behalf of a child conceived on assistance if the entire family had been off aid for two years at some point after the child’s birth or—and here is the important exemption—if the entire family had left cash assistance continuously for at least two months during the ten months prior to the child’s birth. The reference to cash assistance is important; it means that should the family choose to separate from assistance for 60 days, the new child would be exempt from the cap. For a single mother with one child, the cost of this strategy would be two months of benefits in return for the substantially greater long-term benefit of inclusion of the new child on the case. Medi-Cal benefits would not be lost over this period, and the food stamp benefit would increase. In California, families returning to assistance after short departures are not subject to the same requirements that are applied on first entry; in general such “restorations” are fast-tracked. Clearly California’s version of the family cap is much different from the cap as applied in other states, so even if reliable evidence of cap effects on behavior could be obtained from experience elsewhere, its applicability for forecasting the effects of the maximum family grant rule would be questionable.

Nevertheless, it is useful to use the LDB to estimate the number of children of younger ages who would be ineligible for benefit had California's benefit cap policy been in effect in February 1996 and had the cap had no effect save to reduce the grants of families with children under 9 who had been conceived during a period in which the family was receiving assistance. Since this estimate is made at the point in the year in which the LDB and CA237 caseload estimates most closely coincide (see Figure 3) and since data on conception status all involve information recorded earlier, use of 1996 data likely involve little bias, even though the data have not been updated. Table 6 reports the results. In February 1996, 1,006,210 children under 9 were living in California families receiving AFDC cash benefits. Of these, 29.4 percent, or 297,870, were conceived during a period in which their parents were receiving cash assistance.

Suppose for simplicity that each of these children was a second child, so that the marginal benefit figure of $174 (actually, $177 by February 1996) would apply. Of this change, $117 is attributable to the AFDC payment itself; the rest is food stamps. If February may be taken as representative of all 12 months, total savings would amount to 298,000 children multiplied by the marginal benefit of $117 multiplied by 12 months, or $418 million. This is slightly more than 7 percent of total benefit outlays in the state. In 1996, the federal government paid half of California's AFDC costs, so the state share was roughly $209 million. Under TANF, the state gets it all. For recipients, roughly 30 percent of this reduction, if applied, would be offset by an increase in food stamps.

Caution is urged in applying these numbers. No provision is made for the option California provides for women who become pregnant to escape the cap by leaving assistance for two months. In practice some children who meet the definition of "capped" used here would probably be missed, especially if a case were to be opened and closed during the interval between a child's birth and the point

TABLE 6
Welfare Status at Conception and Ethnicity of Children under Age 9 in AFDC Cases in California, February 1996

	AFDC-FG		AFDC-UP		Total	
	Number	Proportion*	Number	Proportion*	Number	Proportion*
Children < 9						
Ethnicity						
Total	766,330	0.30	239,880	0.29	1,006,210	0.30
White (non-Hispanic)	192,070	0.23	57,360	0.27	249,430	0.24
Hispanic	346,170	0.26	124,760	0.22	470,930	0.25
Black	174,220	0.40	11,720	0.42	185,940	0.40
Cambodian	10,450	0.57	2,900	0.59	13,350	0.57
Laotian	10,900	0.60	16,280	0.59	27,180	0.59
Vietnamese	11,410	0.40	13,830	0.47	25,240	0.44
Other	17,670	0.33	10,820	0.47	28,490	0.38
Missing	3,440	0.23	2,430	0.27	5,870	0.25
Cases with Children <9						
Total	517,610	0.33	132,170	0.38	649,780	0.34
White (non-Hispanic)	138,940	0.25	34,260	0.35	173,200	0.27
Hispanic	227,740	0.31	69,240	0.30	296,980	0.31
Black	116,800	0.44	6,140	0.53	122,940	0.44
Cambodian	6,030	0.67	1,510	0.73	7,540	0.68
Laotian	5,310	0.68	5,650	0.77	10,960	0.73
Vietnamese	8,180	0.45	8,590	0.57	16,770	0.51
Other	11,630	0.35	5,370	0.51	17,000	0.40
Missing	2,980	0.18	1,410	0.33	4,390	0.23

Source: Calculated by author using the California Longitudinal Database (UCDATA, 1997). All calculations and estimates are based on a 10 percent sample of cases.

*Estimated proportion of children under 9 conceived when family was receiving cash assistance in California.

of enumeration. These calculations do not include adjustment for the exemption California grants to cases that have been closed for two years (see Table 1). Nevertheless, the amounts are not trivial, especially if they grow over time as the cohort initially subject to the cap ages and as reduced benefits lead to higher termination rates for the families involved.

California, like other states, is enjoying a considerable welfare dividend as the caseload falls but the federal block grant stays fixed. These estimates imply that the dividend is fattened as well by reductions in average benefit per case as the cap prevents the births that routinely appear among families on assistance from increasing benefits. Every year this amount grows as the number of families with one or more children subject to the cap increases.

CONCLUSIONS

This paper has produced the first estimates of the number and cost implications of infants conceived by mothers receiving assistance in California over the period 1988–1995. The estimates are constructed using a longitudinal analytic database derived from administrative records.

The results indicate that such births are a frequent occurrence, but their incidence is declining. In 1988 about 5.2 percent of the AFDC-FG (single-parent) California cases experienced a birth of a child conceived while the case was open; by 1995 this had declined to 4.8 percent. Over the same interval this aggregate rate fell from 8.0 to 6.7 percent for (two-parent) families in AFDC-UP. Nevertheless, the absolute number of such births is large, involving some 62,000 families across both subprograms in 1995.

Under benefit cap policies, children conceived and born while their families were receiving TANF do not produce a change in welfare grant. Though the one-year number of such births looms small in comparison with total numbers of children in AFDC families in 1995, such children accumulate over time, and cases which include such births tend to stay on benefit for a long time. Nearly 30 percent of children under age 9 in families receiving AFDC benefits in February 1996 were conceived while their cases were open; benefits paid on their behalf amounted to roughly 7 percent of total state outlays. The incidence of such births is much higher for some racial/ethnic groups than for others; absent immediate change in birth rates in response to caps, this means that the consequences for family income of the imposition of such policies will differ substantially across groups. The savings resulting from the family cap grow over time, so if the incidence of welfare conceptions in other states is of comparable magnitude to what has been counted here for California, the family cap is contributing substantially to the accelerating decline in welfare costs for states with such policies. The actual effect in California depends upon the extent to which the cap is actually applied and the frequency with which families use the two-month exemption to avoid is effects.

Multivariate analysis indicates that rates of welfare conception are declining. The estimated likelihood that a “representative” recipient mother picked from the AFDC-FG rolls would over the subsequent 12-month interval become pregnant with a child that will be born on welfare declined by 21 percent between 1991 and 1995. In AFDC-UP the decline was greater, but a possible undercount in 1995 throws some doubt on the trend in AFDC-UP. Disaggregation by race supports the downward-trend conclusion for AFDC-FG and leaves the outcome for AFDC-UP ambiguous.

Obviously, some of these issues could be resolved were LDB case extracts for 1997, 1998, and 1999 available. To date such extracts have not been made, apparently due to concerns that the LDB caseload estimates were not tracking those from the CA237 source. The analysis presented here suggests that failure to track reflects transitory factors and that over the long run the LDB has performed well. Given that the effects of the MEDS administration process on the case counts are increasingly well understood, it seems appropriate to resume the extract to look for, among other things, change in welfare conception rates following implementation of the cap in 1997.

Keeping the LDB going is not just a concern of those who like numbers. There are clues for service intervention in these data. Though it will be of little surprise to caseworkers, both the AFDC-FG and AFDC-UP analyses indicate that young women with infants or toddlers who are receiving public assistance today are likely to become pregnant again soon with a child that will be born on assistance. Indeed, for an African-American woman aged 19 with a child 1 year old or less, the multivariate models estimated for this paper imply that the odds of a new pregnancy are 1 in 5 in AFDC-FG and 1 in 4 in the AFDC-UP program. We have no information on whether such pregnancies are desired aside from the fact that they are carried to term. Given the changing terms under which assistance is provided in the world after PRWORA, it is essential that women and their partners be informed of the consequences and the alternatives.

Appendix

Probit Variable Definitions

The population studied includes cases open on February 1 of each year 1988–1995. February 1 is called the point of observation. Cases were eliminated if they included no children at the point of observation or if a child was born to the case prior to November 1 following the point of observation (this restriction eliminates cases in which a woman is already pregnant). The outcome of interest is a conception that commences during the 12 months following the point of observation and is subsequently carried to live birth (births may occur in the subsequent calendar year, and cases may move between counties). Note that some cases may appear in multiple years. Estimation of the models using the Huber-White adjustment to produce robust standard errors and including county fixed effects did not significantly change either estimated coefficients or standard errors, so only uncorrected results are reported in the text. Results for county fixed effects are briefly mentioned below. The value of each indicator variable is conditional; “else 0” is understood.

Variable	**Definition**
INTERCEPT	= 1
Age	Indicator variables for age of oldest female in case on February 1 of sample year. These categories are mutually exclusive.
AGE15–17	= 1 if 15–17
AGE18–19	= 1 if 18–19
AGE20–24	= 1 if 20–24
A6E25–29	= 1 if 25–29
AGE30–34	= 1 if 30–34
AGE35+	= 1 if 35 or older (excluded group)
NOFEMALE	= 1 if case includes no female aged 15+
Race/Ethnicity	Note: The following race/ethnicity categories are derived from coding established by the California Health and Welfare agency. The categories are mutually exclusive. Race/ethnicity was established on the basis of reported value for the oldest female case member aged 15 or greater with race/ethnicity information or, in the absence of females aged $\geq$ 15, the oldest child in the case. This is the “reference person.”
ETHBLACK	= 1 if reference person is black
ETHWHITE	= 1 if reference person is non-Hispanic white (excluded group)
ETHCAMB	= 1 if reference person is Cambodian
ETHHISP	= 1 if reference person is Hispanic
ETHLAOT	= 1 if reference person is Laotian
ETHVIET	= 1 if reference person is Vietnamese

ETHOTHER = 1 if reference person is from group not identified above (includes small number of cases without identified race/ethnicity)

Children Household composition, including indicator for cases at risk of closure due to age of youngest child. Note that categories are not necessarily mutually exclusive.

CHILD=1	= 1 if number of children in case = 1 (excluded group)
CHILD≥2	= 1 if number of children in case ≥ 2
CHILD≥3	= 1 if number of children in case ≥ 3
CHILD≥4	= 1 if number of children in case ≥ 4
YNGSTCHL<2	= 1 if youngest child < 2
YNGSTCHL<3	= 1 if youngest child < 3
YNGSTCHL<6	= 1 if youngest child < 6
YNGSTCHL≥15	= 1 if case includes female aged ≥30 and youngest child aged 15–17

Recent Welfare History Rudimentary information on duration of welfare receipt prior to beginning of year

ONFEB = 1 if case was open in February of preceding year but not open continuously over the succeeding 11-month interval leading to the point of observation.

ONALL = 1 if case was open the entire 12 months preceding the point of observation.

Location

California has 58 counties. To preserve confidentiality of the data, UCDATA aggregated 24 of these into seven rural and/or alpine groups, thus reducing the 58 counties to 41 geographic areas, beginning alphabetically with Alameda County. The AFDC-FG and AFDC-UP probit models were experimentally estimated with indicator variables for the fixed effect, if any, of residence in each county (or county group) on the likelihood of conception subject to the cap. To conserve space, these coefficients are not reported in the text. The results are available from the author. In both equations the number of county effects that were statistically significant ($\alpha = .05$) was larger than would be anticipated by chance. Cases in AFDC-FG in five counties (Kern, Merced, Sacramento, San Joaquin, and Tulare) were more likely to experience a welfare conception than were like cases in Alameda County (the reference group). Cases in five other counties (Orange, Placer, San Luis Obispo, San Mateo, and Yolo) and three of the clusters were less likely. In AFDC-UP cases, families in five counties (Fresno, Merced, Sacramento, San Joaquin, and Sutter/Yuba) and one of the clusters were more likely to exhibit welfare conceptions than were similar families in Alameda County. Families in three counties (Humboldt, Shasta, and Sonoma) were less likely than were their Alameda County counterparts. Note that rates of welfare conception in three counties, Merced, Sacramento, and San Joaquin, are exceptionally high for both subprograms.

Sample Indicators

$Y \geq 1988$	= 1 if observation is for case from February 1, 1988 sample or later (excluded group)
$Y \geq 1989$	= 1 if observation is for case from February 1, 1989 sample or later
$Y \geq 1990$	= 1 if observation is for case from February 1, 1990 sample or later
$Y \geq 1991$	= 1 if observation is for case from February 1, 1991 sample or later
$Y \geq 1992$	= 1 if observation is for case from February 1, 1992 sample or later
$Y \geq 1993$	= 1 if observation is for case from February 1, 1993 sample or later
$Y \geq 1994$	= 1 if observation is for case from February 1, 1994 sample or later
$Y \geq 1995$	= 1 if observation is for case from February 1, 1995 sample or later

References

Acs, Gregory. 1996. “The Impact of Welfare on Young Mothers’ Subsequent Childbearing Decisions.” *Journal of Human Resources* 31: 808–915.

Advisory Panel on Research Uses of Administrative Data. 1998. *Administrative Data for Policy-Relevant Research: Assessment of Current Utility and Recommendations for Development*. Chicago/Evanston Illinois: Northwestern University/University of Chicago Joint Center for Poverty Research.

California Department of Social Services. 1996. *Maximum Family Grant Notice Implementation in the Aid to Families with Dependent Children (AFDC) Program*. All-County Letter No. 96-37. Sacramento.

Camasso, M. J., C. Harvey, R. Jagannathan, M. Killingsworth. 1998a. *A Final Report on the Impact of New Jersey’s Family Development Program: Experimental-Control Group Analysis*. New Brunswick: Rutgers University School of Social Work.

Camasso, M. J., C. Harvey, R. Jagannathan, M. Killingsworth. 1998b. *A Final Report on the Impact of New Jersey’s Family Development Program: Results from a Pre-Post Analysis of Case Heads from 1990–1996*. New Brunswick: Rutgers University School of Social Work.

Committee on Ways and Means. 1998. *1998 Green Book: Background Material and Data on Programs within the Jurisdiction of the Committee on Ways and Means*. Washington, DC: Government Printing Office.

Council of Economic Advisers. 1999. *The Effects of Welfare Policy and the Economic Expansion on Welfare Caseloads: An Update*. Washington, DC.

Fairlie, Robert W., and Rebecca A. London. 1997. “The Effect of Incremental Benefit Levels on Births to AFDC Recipients.” *Journal of Policy Analysis and Management* 16: 575–597.

Horvath, A., and Peters, H. E. 1999. “Welfare Waivers and Nonmarital Childbearing.” Joint Center for Poverty Research. Manuscript.

Loury, Glenn C. 2000. “Preventing Subsequent Births to Welfare Recipients.” American Enterprise Institute. Manuscript.

Moffitt, Robert A. 1998. “The Effect of Welfare on Marriage and Fertility.” In *Welfare, the Family, and Reproductive Behavior*, edited by R. Moffitt. Washington, DC: National Academy Press. Pp. 50–97.

Rossi, Peter. 2000. “New Jersey’s Family Development Program: An Overview and Critique of the Rutgers Evaluation.” Manuscript.

Rowe, Gretchen. 2000. *The Welfare Rules Databook*. Washington, DC: The Urban Institute, forthcoming.

Stark, Shelley, and Jodie Levin-Epstein. 1999. *Excluded Children: Family Cap in a New Era*. Washington, DC: Center for Law and Social Policy.

Turturro, C., B. Benda, and H. Turney. 1997. *Arkansas Welfare waiver Demonstration Project: Final Report*. Little Rock: University of Arkansas.

U.S. Department of Health and Human Services, Office of the Assistant Secretary for Planning and Evaluation. 1997. “Setting the Baseline: A Report on State Welfare Waivers.” Washington, DC.

UCDATA. 1997. “California Work Pays Demonstration Project: Statewide Longitudinal Database—Cases 10% Sample.” 1987–95, Public Use Version 2 Dataset Codebook [machine-readable data file]. Berkeley, California: University of California-Berkeley Data Archive and Technical Assistance.

Ventura, Stephanie J., William D. Mosher, Sally C. Curtin, Joyce C. Abma, and Stanley Henshaw. 2000. *Trends in Pregnancies and Pregnancy Rates by Outcome: Estimates for the United States, 1976–96*. Vital and Health Statistics Series 21, No. 56. DHHS Publication No. (PHS) 2000-1934. Hyattsville, MD: U.S. Department of Health and Human Services, Centers for Disease Control and Prevention, National Center for Health Statistics.

Winn, Ellen, and Mary Clare Lennon. 2000. “The Role of Administrative Data and National Data Sets in Understanding Welfare Reform.” *The Forum* 3(1): 1–6.

Wisconsin Governor’s Office. 1996. *Wisconsin Welfare Reform: A History of Progress, A Plan for the Future*. Madison.

Wiseman, Michael. 1999. “A Management Information Model for New-Style Public Assistance.” Assessing the New Federalism Discussion Paper 99-10. Washington, DC: The Urban Institute, 1999.

Wiseman, Michael. 2000. “Counting the Capped.” In *The Family Cap*, edited by Douglas Besharov and Peter Germanis. College Park, MD: University of Maryland School of Public Policy, forthcoming.