Dynamic Opting-in Incentives in Income-tested Social Programs:

Evidence from Medicaid/CHIP[*]

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October 17 2011

Abstract

Conventional studies of labor supply in the presence of income-tested transfer programs implicitly assume that income eligibility for program participation is constantly monitored by the government. However, this is not how most of these programs operate in practice, and the time until the next eligibility recertification can be as long as a year. In particular, the Balanced Budget Act of 1997 gives states the option of insuring children in their Medicaid/CHIP program continuously for up to 12 months regardless of changes in family income. The long recertification period in effect increases the size of the benefit notch, and neoclassical labor supply models predict that agents may lower their labor supply before the application month to gain program eligibility and then increase their labor supply until the next eligibility check. I use the 2001 and 2004 panels of Survey of Income and Program Participation (SIPP) to empirically examine the income and labor supply responses of parents whose children are publicly insured and find no evidence of the strategic behavior predicted by the neoclassical labor supply models. Comparing calibrated model predictions and the empirical evidence, I can rule out large labor supply elasticities (greater than 0.05 in most comparisons). I discuss the implications of using the length of the continuous eligibility period as a policy instrument, and the empirical evidence suggests that the 12-month period is preferred to those that are shorter.

JEL codes: H21, H24, H51, H53, I18, I38, J22
Keywords: Labor Supply, Medicaid, CHIP, Continuous Eligibility, Recertification

1 Introduction

An implicit assumption in labor supply studies of income-tested transfer programs is that program eligibility is being constantly monitore. However, this is not how these programs operate in reality and the time between two eligibility certifications can be as long as a year. Although this fact is recognized in several economics studies[1], a formal investigation has not been carried out to address how program participants adjust their labor supply behavior in response to the dynamic incentives created by the lack of constant income monitoring. In this paper, I attempt to fill this gap by examining families' behavioral responses to the continuous eligibility provision in Medicaid and State Children's Health Insurance Program (SCHIP or simply CHIP) for children.

Intuitively, there are two channels through which a program becomes more valuable to participants when the period of continuous benefits are increased. First, transaction costs associated with renewal of benefits decrease because of less frequent renewal periods, as pointed out by Currie et al. (2001) and Kabbani and Wilde (2003). The second, and rarely considered point, is that the guarantee of continuous eligibility effectively changes a participant's budget constraint. If the budget constraint non-linearity created by the eligibility requirements distort a family's labor supply choice, any period in which eligibility is not checked will eliminate the distortion and allow the family a more optimal consumption bundle. Increasing the re-certification period effectively decreases the number of periods that a family will face the more stringent budget constraint, creating strong incentives for an otherwise ineligible family to opt into the program. That is, families may be induced to temporarily lower their income, gain program eligibility, and revert back to their "optimal" consumption bundle after having acquired the government benefit for the entire continuous eligibility period.

As the length of the continuous eligibility period increases, more families may be expected to participate in the program, resulting in increased expenditure and downward budget pressure. At the same time, these newly participating families of higher income are arguably not the intended beneficiaries of the government transfer. On the other hand, if the continuous eligibility period is significantly shortened, the transaction costs of frequent eligibility recertifications may become insurmountable for a part of the low-income group who may be most in need of the transfer (e.g. single parent families–see Currie et al. (2001)). In the case of

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[1]In the context of Food Stamp, for example, Currie et al. (2001) and Kabbani and Wilde (2003) find that shortening the recertification period reduces participation rate. The two surveys Currie and Gahvari (2008) and Currie (2004) both mentioned (potentially) long recertification periods in transfer programs, and the latter noted that WIC has a fixed recertification time during which families are eligible irrespective of income changes.

health insurance, studies (e.g. Olson et al. (2005)) have shown that children who experience interruptions in health insurance coverage are more likely to have unmet health care needs, and therefore imposing large transaction costs on otherwise eligible families is socially suboptimal. Furthermore, verifying eligibility in short intervals leads to increases in administrative cost for the government as well.

Given the tradeoffs of increasing the continuous eligibility period, understanding the behavioral reaction of economic agents to the lack of monitoring has important policy implications. If extensive strategic dip-and-rebound behavior in income is found, then it may suggest that the recertification period is too long. If no strategic behavior is found, on the other hand, marginally increasing the continuous eligibility period may be desirable. I will formally present these arguments in a later part of this paper.

After introducing the theoretical model and predictions, I investigate the empirical implications in the context of Medicaid/CHIP. Along with creating the SCHIP program, the Balanced Budget Act of 1997 gave states the option of continuously insuring children for up to 12 months in their public insurance programs regardless of changes in family income during that period. As a result, a third of the states implemented the continuous eligibility option in their public insurance program for children. These states present an opportunity to gauge the significance of the strategic behavior, which sheds light on the choice of the optimal continuous eligibility period.

The contributions of this paper are as follows. First, I recognize the potential dynamic impact of a long continuous eligibility or recertificatin period on the labor supply decisions of program participants. I derive qualitative and quantitative predictions of the family income process using neo-classical labor supply models that incorporate the budget constraint relevant for continuous eligibility. Second, I empirically examine the model predictions using SIPP data. Third, I discuss the implication of using the continuous eligibility period as a policy instrument in light of the empirical finding.

Empirically, I find no evidence of the short-term dip-and-rebound strategic behavior in average income as predicted by the neoclassical models. Comparing calibrated model predictions and the empirical evidence points to a small labor supply elasticity (less than 0.05 in most comparisons), which is consistent with the estimates in studies using nonlinearities in the tax schedules (e.g. Saez (2010) and Chetty et al. (2011)). Following a formal analysis of the optimal continuous eligibility period, I conclude that the lack of strategic behavior suggests that it may be optimal for states that still have a 6-month renewal period, namely Georgia and Texas, to consider halving the renewal frequency; it may also be optimal for those currently offering 12 months of continuous eligibility not to switch back to a 6-month period as in the case of–for example–

Connecticut, Indiana, Nebraska, Washington and New Mexico in the early 2000's.

The remainder of the paper is organized as follows. Section (2) provides an overview of the Medicaid/CHIP institutions. Section (3) presents a series of labor supply models to theoretically analyze families' responses to the continuous eligibility provision. Section (4) describes data used, and empirical results are presented in Section (5). Section (6) calibrates the labor supply model and compares the quantitative prediction to the empirical results. Section (7) discusses the implication of the empirical findings on the optimal choice of the continuous eligibility period length. Section (8) concludes.

2 Institutional Background of Medicaid and CHIP

The Medicaid program was created by the Social Security Amendments of 1965 and provides health insurance to low-income populations. The program originally targeted those traditionally eligible for welfare–single-parent families, and the aged, blind and disabled. However, eligibility for public insurance through Medicaid, and later, through SCHIP, has expanded substantially over time particularly for the population of dependent children.

Over the 1980s, the link between Medicaid and welfare for children was gradually severed through a series of legislative acts. In 1984, the Deficit Reduction Act required states to cover children less than five years old born after September 30, 1983 living in families income-eligible for Aid to Families with Dependent Children (AFDC), regardless of family structure. Further decoupling occurred with passage of the Omnibus Budget Reconciliation Acts (OBRA) of 1986 and 1987, which allowed states to raise the income limits for Medicaid eligibility above the AFDC thresholds. OBRA 1987 also required states to cover all children less than seven years old born after September 30, 1983 living in families with incomes below the AFDC income threshold. Pregnant women and infants living in families with incomes below 75% of the federal poverty level (FPL) were granted mandatory eligibility through the Medicare Catastrophic Coverage Act of 1988. Also in 1988, the passage of the Family Support Act required states to continue Medicaid coverage for up to one year for families who lost AFDC benefits due to increased earnings.

The two largest federal expansions were included in OBRA 1989 and OBRA 1990, which became effective in April 1990 and July 1991 respectively. OBRA 1989 required states to offer Medicaid coverage to pregnant women and children up to age six with family incomes below 133% of the FPL. OBRA 1990 required states to cover children born after September 30, 1983 with family incomes below 100% of the

FPL. The two expansions remain the mandated minimum federal standards for children today: a child under the age 6 is eligible for Medicaid if her family income is below 133% of the FPL, and a child between the age of 6 and 18 is eligible if her family income is below 100% of the FPL (for a detailed account of the major Medicaid legislations by 1997, see Gruber (2003)).

While states are required to adhere to these minimum federal standards from OBRA 1990, the creation of the State Children's Health Insurance Program (SCHIP) in 1997 allowed many states to further expand their public insurance programs above these standards. Unlike Medicaid, SCHIP provided states with block grants to fund coverage for children and left the implementation of program up to the individual states, subject to some rules to prevent crowd out of private insurance and to meet federal benefit standards. Specifically, states could choose to use their funds by expanding their existing Medicaid program, creating a separate program for children who do not qualify for the existing Medicaid program, or a combination of both. In some separate state programs, families who exceed a certain income level were also required to pay a premium for coverage. As a result of the block granting structure of SCHIP, states varied widely in their implementation of public insurance for children. One particular feature of some state programs is a continuous eligibility period as permitted by the Balanced Budget Act of 1997, which provided children with uninterrupted coverage for up to 12 months after confirming eligibility, regardless of whether their families' incomes rise above the eligibility requirement during this period. The focus of this paper is the effect of this continuous eligibility period on public insurance coverage for children and on the labor supply of their families.

In practice, Medicaid and CHIP eligibility is established based on the most recent monthly income, and official income proof[2] needs to be submitted with the application in most cases. However, with the implementation of a continuous eligibility period, some states explicitly mandate that once a family is qualified for coverage, they are eligible for coverage for a fixed, continuous period after that point. After this period, the family is once again required to confirm their eligibility, either by reporting any changes in income, or by actually sending proof of income with a renewal application.

Finally, along with the continuous eligibility provision, the Balanced Budget Act of 1997 also gave states the option of allowing presumptive eligibility for children. That is, states may allow children who

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[2]On the current application form for New York State, for example, it states that an applicant must provide a letter,written statement, or copy of check or stubs, from the employer, person or agency providing the income... [the applicant should] [p]rovide the most recent proof of income before taxes and any other deductions. The proof must be dated, include the employee's name and show gross income for the pay period. The proof must be for the last four weeks, whether you get paid weekly, bi-weekly, or monthly. It is important that these be current".

appear eligible to obtain temporary Medicaid/CHIP eligibility (so that they may immediately access health care services) while their eligibility based on income are being confirmed.[3] Since children who are covered under presumptive eligibility do not always need to meet the usual income requirements, I analyze the effect of the continuous eligibility provision on family income both including and excluding public insurance spells that are due to presumptive eligibility. Twelve states provided presumptive eligibility to children in my sample and are listed in the Appendix.

3 Theoretical Framework

This section derives the predictions of various economic models regarding the families' response to the continuous eligibility provision in Medicaid/CHIP. Subsection (3.1) first reviews the prediction of a standard static neo-classical labor supply model in the presence of an in-kind transfer, and I show in subsection (3.2) that the dynamic problem with continuous eligibility provisions can be reduced to two static problems with different budget constraints. A solution is provided for the dynamic model where the flow utility is quasi-linear, and it predicts a dip and rebound in average income at each eligibility check.

3.1 Constant Eligibility Recertification: Baseline Static Model

In this section, I analyze the labor supply decisions when eligibility for Medicaid/CHIP is constantly recertified. That is, families are eligible for benefits only if their income is below a cutoff as is assumed in the conventional labor supply framework in the literature. The analysis is standard, and implications–at least those qualitative–have been explored in other studies (e.g. Blank (1989) and Yelowitz (1995)), but I will illustrate it here using the functional form from Saez (2010) because results derived below will be relevant for the dynamic problem with continuous eligibility provision in section (3.2). In addition, subsection (3.1.2) examines the implication of allowing agents only discrete labor supply choices as opposed to giving them the freedom to perfectly control their income. Subsection (3.1.3) explores the consequence of introducing welfare stigma, income effect and allowing heterogeneity in the elasticity of labor supply.

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[3]Presumptive eligibility to infants and pregnant women were granted a decade earlier by OBRA '86.

3.1.1 Continuous Labor Supply Choice

The function form of the baseline model is taken from Saez (2010), which studies the bunching behavior of economic agents in response to kinks in the tax schedule. The particular utility functional form has also been used in other recent papers, e.g. Chetty et al. (2011), that study the response to nonlinearities in the budget constraint.

Agents[4] choose *continuous* labor supply H and consumption C to maximize utility that increases in C and decreases H. Specifically, the utility function is of quasi-linear form

$$u(C,H) = C - \frac{n}{1+1/e}(\frac{wH}{n})^{1+1/e} \tag{1}$$

where w is wage rate, and n and e are parameters indicating taste and the responsiveness of pre-tax income to a change in tax rate.[5] Solving the optimization problem implies that an agent chooses optimal labor supply $H^* = \frac{n}{w}(1-t)^e$ and hence optimal pre-tax income $wH^* = n(1-t)^e$ when facing the budget constraint $C = (1-t)wH$. As pointed out by Saez (2010), $wH^* = n$ when $t = 0$, and n can be interpreted as the choice of potential income in the absence of a marginal tax. An agent with a larger n both work and consume more, and n is assumed to be smoothly distributed according to density f_n across the population.

e is the elasticity of labor supply with respect to (one minus) the marginal tax rate because of the following identity $\frac{(1-t)}{(H^*)}\frac{d(H^*)}{d(1-t)} = e$. As is well known, the quasi-linear utility functional form implies no income effect, so e is both the compensated and uncompensated elasticity. I assume e to be constant in this section but the consequence of allowing heterogeneity in e, as well as allowing income effect, will be discussed in section (3.1.3).

The presence of Medicaid/CHIP induces at least one notch [6] in the budget constraint. For simplicity of exposition, I include only one notch and a single marginal tax rate in the presentation below, and the notch is the benefit received for a family whose children just qualify for CHIP. The empirical impacts of other notches (induced by Medicaid or other transfer programs) and differential tax rates will be discussed

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[4]The word "agent" will be used interchangeably with "family" in this section, and the decision of an agent encapsulates that of parents in a family.

[5]The arguments of the utility function in the Saez model are pre-tax income Z and post-tax income C where the disutility of Z originating from working is implicit. Here Z is written explicitly as wH where w is considered to be distributed smoothly among agents. I make this change for the analyses carried out in section (3.1.2).

[6]As mentioned in section (2), families enrolled in CHIP with income above the 150% FPL may be subject to moderate premiums and co-payments, which implies a lower CHIP notch than that of Medicaid.

in section (6). Formally, the budget constraint a family faces is

$$C = [wH(1-t) + g]1_{[wH \leqslant \gamma]} + wH(1-t)1_{\{wH > \gamma\}} \tag{2}$$

where γ is the Medicaid/CHIP eligibility cutoff, g the monthly value of public insurance and t the marginal tax rate. As pointed out by Blank (1989), no family will choose income to be just above the threshold. This is intuitive because a family consumes more and works less by choosing its income to be at the eligibility cutoff than just above it. Certain families who would have chosen $H > \frac{\gamma}{w}$ in the absence of Medicaid/CHIP would now switch to γ. Solving the optimization problem predicts the choice of H for a family of type n:

$$H = \begin{cases} \frac{n}{w}(1-t)^e & \text{if } n \leqslant n_\gamma \text{ or } n > \bar{n} \\ \frac{\gamma}{w} & \text{if } n \in (n_\gamma, \bar{n}) \end{cases}$$

where $n_\gamma = \frac{\gamma}{(1-t)^e}$ is the type of agent who choose income at γ in the absence of the notch and $\bar{n}$ is the highest type of agent who would bunch at γ in the presence of the notch. An agent with $\bar{n}$ is indifferent between the consumption-work bundle at the notch $(\gamma(1-t) + g, \frac{\gamma}{w})$ and her optimal choice in the absence of notch $(\bar{n}(1-t)^{1+e}, \frac{\bar{n}(1-t)^e}{w})$. Therefore, $\bar{n}$ is the solution to the equation

$$\gamma(1-t) + g - \frac{\bar{n}}{1+1/e}(\frac{\gamma}{\bar{n}})^{1+1/e} \;=\; \bar{n}(1-t)^{1+e} - \frac{\bar{n}}{1+1/e}(1-t)^{1+e} \tag{3}$$

It follows that the distribution of pre-tax income $Z = wH^*$ is given by:

$$f_Z(z) = \begin{cases} \frac{1}{(1-t)^e} f_n(\frac{z}{(1-t)^e}) & \text{if } z < \gamma \text{ and } z \geqslant \bar{n}(1-t)^e \\ 0 & \text{if } z \in (\gamma, \bar{n}(1-t)^e) \end{cases}$$

and

$$\Pr(Z = \gamma) = F_n(\bar{n}) - F_n(n_\gamma)$$

In summary, the standard static model makes the prediction that, when a benefit notch is introduced, agents originally choosing income just above the eligibility cutoff will lower their labor supply to locate at the cutoff and become just eligible for benefit. Thus, there is income bunching at the eligibility cutoff and a drop in density to 0 right above the cutoff.

3.1.2 Discrete Labor Supply Choices

In the preceding section, agents' labor supply choice is assumed to be continuous which implies that agents are free to choose their hours and hence perfectly control their income. Obviously, this may not be a realistic restriction per Ashenfelter (1980), Ham (1982), Kahn and Lang (1991), Altonji and Paxson (1992), Dickens and Lundberg (1993) and Chetty et al. (2011), to name a few. In this subsection, I will first derive the theoretical prediction by only allowing an agent finitely many hours choices.[7] The main implication is still that certain agents will lower their labor supply in order to claim benefit when a notch is introduced. But rather than bunching at the eligibility cutoff, there is only a discontinuous drop in the density of income at the cutoff.

For exposition purposes, I will discuss in this section the case when H can only vary along the extensive margin; that is, an agent can only work full time or not work at all. The general case where H is allowed more than two choices is explored in the Appendix. Let $H = 0$, $H = 1$ denote the labor supply choice of not working and working full time respectively. If workers are constrained to only these two labor supply options, then the maximization problem becomes $\max_{H \in \{0,1\}} u(C,H)$ subject to the budget constraint (2), and we solve the maximization problem by considering the following two scenarios.

1. $w \leqslant \gamma$. For an agent with potential monthly wage below the cutoff, she can claim benefit whether she works or not. In other words, the budget constraint they face is only the segment to the left of γ: $C = (1-t)wH + g$. Consequently, maximizing utility involves the comparison of $u(g,0)$ and $u((1-t)w+g,1)$. To characterize the solutions, consider the agent of type $\bar{n}^{l}$[8] who is indifferent between choosing $H = 0$ and $H = 1$ at wage w. Therefore, $\bar{n}^{l}$ solves

$$u(g,0) \equiv g = (1-t)w + g - \frac{n}{1+1/e}\left(\frac{w}{n}\right)^{1+1/e} \equiv u((1-t)w+g,1)$$

which implies that $\bar{n}^{l}(w) = \left(\frac{w^{1+1/e}}{[(1-t)w](1+1/e)}\right)^{e}$. Since $\frac{n}{1+1/e}\left(\frac{w}{n}\right)^{1+1/e}$ is decreasing in n, i.e. the disutility of working is less for an agent with high n, agents with wage w and of type $n \geqslant \bar{n}^{l}(w)$ choose $H = 1$ and those with $n < \bar{n}^{l}(w)$ choose $H = 0$.

2. $w > \gamma$. For an agent with potential monthly wage above the cutoff, she is eligible for benefits only if she chooses not to work. The type of agent who is indifferent between working and not working at wage w

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[7] The working paper versions of Saez (2010), Saez (1999) and Saez (2002), address this extension in their simulation section but do not discuss the predictions from a theoretical perspective.

[8] The superscript l here stands for low wage. h will be used for the next case.

equates $u(g,0)$ and $u((1-t)w,1)$. Because her type $\bar{n}^r$ solves

$$u(g,0) \quad \equiv \quad g = (1-t)w - \frac{n}{1+1/e}(\frac{w}{n})^{1+1/e} \equiv u((1-t)w,1)$$

$\bar{n}^r(w) = (\frac{w^{1+1/e}}{[(1-t)w-g](1+1/e)})^e.$[9] Analogous to the case above, agents with $n \geqslant \bar{n}^r(w)$ choose to work full time while those with $n < \bar{n}^r(w)$ choose not to work.

To summarize, if $\bar{n}_{0,1}$ denotes the type of agents who are indifferent between working and not working, then

$$\bar{n}_{0,1}(w) = \begin{cases} \bar{n}^l(w) & \text{if } w \leqslant \gamma \\[2mm] \bar{n}^r(w) & \text{if } w > \gamma \end{cases}$$

$\bar{n}_{0,1}$ varies smoothly with w within each case, but there is a discontinuous increase in $\bar{n}_{0,1}$ as w crosses γ. When w and n follow a smooth joint distribution $f_{n,w}$ over the first quadrant of $\mathbb{R}^2$, this discontinuous drop in threshold agent type implies no bunching but a discontinuity in the density of pre-tax income $Z = wH$ at γ. To see this, notice that the c.d.f of Z evaluated at $z > 0$ is

$$F_Z(z) \quad = \quad \Pr(Z=0) + \Pr(0 < Z \leqslant z) = \Pr(H=0) + \Pr(H=1, w \leqslant z) \tag{4}$$

On two sides of the eligibility cutoff γ, the values of F_Z are

$$F_Z(z) \quad = \quad \begin{cases} \Pr(H=0) + \Pr(w \leqslant z, n \geqslant \bar{n}^l(w)) & \text{if } z \leqslant \gamma \\[2mm] \Pr(H=0) + \Pr(w \leqslant \gamma, n \geqslant \bar{n}^l(w)) + \Pr(\gamma < w \leqslant z, n \geqslant \bar{n}^r(w)) & \text{if } z > \gamma \end{cases}$$

$$= \quad \begin{cases} \Pr(H=0) + \int_0^z \int_{\bar{n}^l(w')}^\infty f_{n,w}(n',w') dn' dw' & \text{if } z \leqslant \gamma \\[2mm] \Pr(H=0) + \int_0^\gamma \int_{\bar{n}^l(w')}^\infty f_{n,w}(n',w') dn' dw' + \int_\gamma^z \int_{\bar{n}^r(w')}^\infty f_{n,w}(n',w') dn' dw' & \text{if } z > \gamma \end{cases}$$

Since $\int_0^z \int_{\bar{n}_l(w')}^\infty f_{n,w}(n',w') dn' dw'$ is continuous in z and that $\lim_{z \downarrow \gamma} \int_\gamma^z \int_{\bar{n}^r(w')}^\infty f_{n,w}(n',w') dn' dw' = 0$, $F_Z(z)$ is continuous at γ. Hence, there is no bunching at the eligibility cutoff unlike in section (3.1.1) where agents can choose along the intensive margin of labor supply.

However, the p.d.f of Z, $f_Z(z)$, is not continuous at γ. By continuity of $f_{n,w}$, $\bar{n}^l$ and $\bar{n}^r$ along with an

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[9]Note that a positive n^r exists–n^r has to be positive for the marginal utility of work to be negative–when $(1-t)w > g$, which means that the post-tax income of working full time at wage w is larger than the value of benefit g. This is most likely satisfied in reality for families with a wage above the CHIP cutoff.

application of the Fundamental Theorem of Calculus,

$$\lim_{z \uparrow \gamma} f_Z(z) = \int_{\bar{n}^l(\gamma)}^{\infty} f_{n,w}(n', \gamma) dn'$$

$$\lim_{z \downarrow \gamma} f_Z(z) = \int_{\bar{n}^r(\gamma)}^{\infty} f_{n,w}(n', \gamma) dn'$$

Since $\bar{n}^r(\gamma) > \bar{n}^l(\gamma)$, $\lim_{z \uparrow \gamma} f_Z(z) > \lim_{z \downarrow \gamma} f_Z(z)$ which implies a discontinuous drop in the income density at γ.

The constraint that workers can only work full time or not work at all is too restrictive. In reality, workers may and do work part time. It is plausible that employers offer several hours-of-work choices to their employees. As shown in the Appendix, the result of no bunching but a density discontinuity at cutoff holds true when H takes on a finite number of values.

3.1.3 Heterogeneous Elasticity, Welfare Stigma and Income Effect

In the previous subsections (3.1.1) and (3.1.2), the labor supply elasticity e is held constant across agents, perfect compliance (i.e. those eligible will participate in Medicaid/CHIP) is assumed, and there are no income effects in the utility function (1). This subsection investigates the implication of relaxing these assumptions and shows that the qualitative predictions in the previous subsections still hold true.

1. Heterogeneous labor supply elasticity. Instead of requiring agents to share the same labor supply elasticity, the first part of this subsection studies the pre-tax income distribution when elasticity is heterogeneous across families. In both sections (3.1.1) and (3.1.2), the threshold taste parameters, i.e. the $\bar{n}$'s, are a function of e, and all statements are true for each $e > 0$. Now suppose that e is heterogeneous and distributed smoothly across agents. In the case where there is no constraint on labor supply, the discontinuous drop in the income density at γ is

$$\lim_{z \uparrow \gamma} f_Z(z) - \lim_{z \downarrow \gamma} f_Z(z) = \int_0^{\infty} \frac{1}{(1-t)^e} f_{n|e}(n_\gamma(e)|e) f_e(e) de \tag{5}$$

[10] and the fraction of agents bunching at γ is

$$\Pr(Z = \gamma) = \int_0^{\infty} (F_{n|e}(\bar{n}(e)|e) - F_{n|e}(n_\gamma(e))) f_e(e) de \tag{6}$$

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[10]Note that $\lim_{z \uparrow \gamma} f_Z(z) = \int_0^{\infty} \frac{1}{(1-t)^e} f_{n|e}(n_\gamma|e) f_e(e) de$ and $\lim_{z \downarrow \gamma} f_Z(z) = 0$.

where $\bar{n}(e)$ and $n_\gamma(e)$ are as defined in subsection (3.1.1)–$\bar{n}(e)$ is the solution to (3) and $n_\gamma(e) = \frac{\gamma}{(1-t)^e}$. Since the integrand in both (5) and (6) is positive, there is a still bunching and a discontinuous drop in the income density at γ. Analogously, the result of no bunching but a density discontinuity at γ from subsection (3.1.2) also holds when labor supply is constrained to several choices–since the result holds for all e, it also holds when integrating over the density of e.

2. Non-participation. To account for non-participation among eligible agents, the literature resorts to welfare stigma proposed by Moffitt (1983). Despite its name, the stigma term can encapsulate much more than the simple psychological cost of being perceived as a beneficiary of government programs. For example, it may incorporate the fixed cost of applying for benefit such as filling out the required forms and learning about program rules. The simplest formulation of welfare stigma is a flat cost to participating in welfare programs. The maximization problem becomes:

$$\max_{C,Z,P} u(C,Z) - \phi P$$

where agents' welfare participation decision $P \in \{0,1\}$ depends on the stigma parameter $\phi > 0$.

In effect, introducing welfare stigma shifts down the program segment of the budget constraint $[wH(1-t)+g]1_{[wH\leqslant\gamma]}$ by ϕ and therefore reduces the public insurance notch to $\max\{g-\phi,0\}$. If ϕ is constant across agents and $\phi < g$, then all the analyses in (3.1.1) and (3.1.2) carry through by replacing g with $\tilde{g} = g - \phi$. When ϕ is heterogeneous, the income distribution is smooth for the sub-population with $\tilde{g} = g - \phi \leqslant 0$, and analyses from previous subsections only hold true for the those with $\tilde{g} > 0$. In the entire population, the qualitative predictions from (3.1.1) and (3.1.2) are still valid if (n,w,e,ϕ) follows a smooth distribution supported on $\mathbb{R}^4_{++}$, although the bunching and density discontinuities are less pronounced due to the existence of non-participants.[11]

3. Non-zero income effects. As mentioned in section (3.1.1), the quasi-linear functional form of (1) eliminates the income effects. This may be reasonable in the context of a tax rate change (i.e. a kink in the budget constraint) as Chetty et al. (2011) stated "tax rates at kinks that [they] exploit for identification has little effect on average tax rates and thus generates negligible income effects". In the case of a notch, however, the absence of income effects in modeling may no longer be appropriate. Here I explore the

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[11]Note that allowing heterogeneity in ϕ is equivalent to allowing heterogeneity in g, but broader interpretation of the heterogeneity in the notch size is permitted. For example, families with healthier children arguably value health insurance less than those with sicker children and would hence face a larger notch.

implication of using a functional form that allows non-zero income effects.

Consider the utility function

$$u(C,H) = \frac{C^{1-\rho}}{1-\rho} - \frac{n}{1+1/e}(\frac{wH}{n})^{1+1/e} \tag{7}$$

whose consumption part displays constant relative risk version, and it encapsulates the quasi-linear utility (1) as a special case when $\rho = 0$. When facing a budget constraint $C = (1-t)wH$, the optimal interior labor supply choice is $H^* = \frac{1}{w}(1-t)^{\frac{e}{\rho e+1}}n^{\frac{1}{1+\rho e}}$ which implies optimal pre-tax income $Z = wH^* = (1-t)^{\frac{e}{\rho e+1}}n^{\frac{1}{1+\rho e}}$. Therefore, $n^{\frac{1}{1+\rho e}}$ is the agent's desired income choice when $t = 0$. The elasticity of labor supply with respect to marginal tax rate becomes $\frac{\partial H^*}{\partial(1-t)}\frac{1-t}{H^*} = \frac{e}{\rho e+1} < e$ when $\rho > 0$. Intuitively, income effects render workers consume more leisure as tax rates go down and consequently labor supply is not as responsive to a change in marginal tax rates.

The analyses undertaken in sections (3.1.1) and (3.1.2) carry through with the more general utility function (7) although the expressions for the various $\bar{n}$'s will change. Therefore, the introduction of non-zero income effects does not change the qualitative predictions. That is, there is income bunching at the eligibility cutoff when agents have prefect control over their income and a discontinuous drop in income density at the cutoff when agents face a menu of finitely many labor supply choices. The intuition is that these predictions hinge on the convexity of the indifference curves, which is not altered when curvature in consumption utility is introduced.

3.2 Continuous Eligibility–Dynamic Models

This section extends the static framework in the previous section to incorporate continuous eligibility provisions. In essence, the provisions allow a more generous budget constraint over time than (2). More specifically, families that are just approved for public insurance can have income above γ and remain covered until the eligibility recertification a year later. To characterize a family's consumption and labor supply decisions in the presence of continuous eligibility provisions, I cast the family's utility maximization problem in a dynamic programming framework.

Formally, the state variable s is the number of months until recertification (s is defined to be 0 for those not claiming benefit since they will face eligibility check when they apply), and let τ be the number of months provided continuous eligibility. In each period, an agent chooses whether or not to participate in the

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program:

$$V_s = \max_{P_s} P_s V_s^1 + (1 - P_s)V_s^0$$

where $P_s = 0, 1$ denotes participation choice, and V_s^1 and V_s^0 are utilities associated with participating and not participating in the program when agents are s months away from eligibility check. Formally, the expressions for V_s^1 and V_s^0 are

$$V_s^1 = \max_{C,Z}\{u(C,H) + \beta V_{s'}\}$$

s.t. $wH < \gamma$ if $s = 0$; $C = (1-t)wH + g$

$$s' = \begin{cases} s-1 & \text{if } s > 0 \\ \tau - 1 & \text{if } s = 0 \end{cases}$$

$$V_s^0 = \max_{c,z}\{u(C,H) + \beta V_{s'}\}$$

s.t. $C = (1-t)wH$

$$s' = \begin{cases} s-1 & \text{if } s > 0 \\ 0 & \text{if } s = 0 \end{cases}$$

For illustration purposes, first consider the simple case when $\tau = 2$, in which case s takes on the value 0 or 1. Let $\{C_s^p, H_s^p\} = \text{argmax}V_s^p$ for $p = 0, 1$. The dynamic problem is thus simplified to

$$V_0 = \max_{P_0} P_0\{u(C_0^1, H_0^1) + \beta V_1\} + (1 - P_0)\{u(C_0^0, H_0^0) + \beta V_0\}$$

$$V_1 = \max_{P_1} P_1\{u(C_1^1, H_1^1) + \beta V_0\} + (1 - P_1)\{u(C_1^0, H_1^0) + \beta V_0\}$$

and I will characterize the optimal P_s, C_s^p and H_s^p's below.

First note that choosing $P_1 = 1$ strictly dominates $P_1 = 0$ because (C_1^0, H_1^0) lies in the interior of the budget set for an agent with $s = 1$. In other words, when benefit can be claimed at no cost (i.e. no restrictions on income), a rational family will do so. This reasoning simplifies the expression for V_1: $V_1 = u(C_1^1, H_1^1) + \beta V_0$. Plugging in this expression of V_1 into that of V_0 leads to

$$V_0 = \max_{P_0} P_0\{u(C_0^1, H_0^1) + \beta u(C_1^1, H_1^1) + \beta^2 V_0\} + (1 - P_0)\{u(C_0^0, H_0^0) + \beta V_0\}$$

For the agents indifferent between choosing $P_0 = 0$ and $P_0 = 1$,

$$V_0 = u(C_0^1, H_0^1) + \beta u(C_1^1, H_1^1) + \beta^2 V_0 = u(C_0^0, H_0^0) + \beta V_0$$

and therefore $V_0 = \frac{u(C_0^0, H_0^0)}{1-\beta}$. It follows that

$$u(C_0^1, H_0^1) + \beta u(C_1^1, H_1^1) = u(C_0^0, H_0^0) + \beta u(C_0^0, H_0^0) \tag{8}$$

If u is functional form in (1), then $C_1^1 = C_0^0 + g$ and $H_1^1 = H_0^0$ because of quasilinearity. Consequently, $u(C_1^1, H_1^1) = u(C_0^0, H_0^0) + g$, and (8) leads to

$$u(C_0^1, H_0^1) + \beta g = u(C_0^0, H_0^0) \tag{9}$$

Note that the (C_0^1, H_0^1) that satisfies (9) has to be a corner solution with $wH_0^1 = \gamma$, for otherwise $u(C_0^1, H_0^1) + \beta g > u(C_0^0, H_0^0)$. Denote the indifferent agent's type by $\bar{n}^d$ and expanding (9) using the quasi-linear functional form leads to

$$\gamma(1-t) + (1+\beta)g - \frac{\bar{n}^{dynamic}}{1+1/e}(\frac{\gamma}{\bar{n}^{dynamic}})^{1+1/e} = \bar{n}^{dynamic}(1-t)^{1+e} - \frac{\bar{n}^{dynamic}}{1+1/e}(1-t)^{1+e} \tag{10}$$

Equation (10) states that an agent of type $\bar{n}^{dynamic}$ is indifferent between choosing her interior solution on the budget constraint segment $C = (1-t)wH1_{\{wH>\gamma\}}$ and the consumption-leisure bundle $(\gamma(1-t) + (1+\beta)g, \frac{\gamma}{w})$. Analogous to the analyses in section (3.1.1), agents with $n \leqslant \bar{n}^{dynamic}$ choose to participate in the program and those with $n > \bar{n}^{dynamic}$ do not. While those with $n \in (0, n_\gamma]$ participate without altering their supplies, those with $n \in (n_\gamma, \bar{n}^{dynamic}]$ will lower their labor supply and income at $s = 0$ to gain eligibility but revert back to their desired interior solution with an income above γ when their eligibility is not checked (i.e. $s > 0$). Comparing (10) to (3) reveals that doubling the length of the recertification period in effect doubles the benefit notch if $\beta \approx 1$. It is easy to show that for a general recertification period τ, the size of the benefit notch an agent faces is effectively $\sum_{i=0}^{\tau-1} \beta^\tau g \approx \tau g$ when making her participation decision at $s = 0$.

There are two extensions to consider. First if a cost is associated with applying for benefit, then the marginal agents at period $s = 0$ will compare choosing a notch of size $\tau g - \phi$ to their interior labor supply choice on $C = (1-t)wH$ where ϕ entails the application cost. The results above hold for agents whose $\phi < \tau g$. Second, when income effects are take into account, the solution may no longer be obtained analytically. Qualitatively, once a new applicant family is approved for benefit, the transfer may reduce labor supply through the channel of income effect. This will imply that the rebound in income after starting a public insurance spell will not be as large as when income effect is absent. In section (6), I will calibrate the model

that allows income effect and compare the predicted effect to that observed empirically.

To summarize, the dynamic model with quasi-linear flow utility and continuous hours choices makes the following predictions:

1. Once eligible, agents will remain in the program.

2. Average income drops at eligibility, and the size of the drop and rebound depends on the income effects (ρ).

In the following sections, I will examine whether agents empirically behave as predicted by the labor supply model.

4   Data and the Construction of the Analysis Sample

To examine the income and labor supply responses to the continuous eligibility provisions in Medicaid/CHIP, I use data from the 2001 and 2004 panels of the Survey of Income and Program Participation (SIPP). SIPP is a representative household survey designed to provide detailed information on incomes and labor force and government program participation. Each of the panel files contains four rotation groups, and they span the period from Oct 2000 to Dec 2003 and Oct 2003 and Dec 2007 respectively. All of the rotation groups in the 2001 panel provide information for 36 consecutive months, and those in the 2004 panel for 48 months. Each adult member of the participating household was interviewed every four months about his or her experiences since the last interview (i.e. four-month reference period).

The chief advantage of SIPP over other candidate datasets (CPS, PSID, HIS, etc) is its panel structure at the monthly frequency and the rich array of variables including detailed information on income, program participation, and family structures. Since the focus of the study is to examine families' income and labor supply behavior before and during their children's Medicaid/CHIP spells, SIPP is the best choice among public use survey data sets for the purpose of this study.

There are, however, several limitations of the SIPP data. First is the existence of the well-known seam bias, which refers to the fact that changes in status are underreported within a four-month reference period while overreported between two reference periods (see, for example, Ham et al. (2009) for details). As noted above, the interviews are not conducted every month but every four months, and children's reproted public insurance coverage are therefore much more likely to start on the first month of the reference period than the second, third or fourth. In fact, about 80% of the fresh spells in the 2001 panel and 90% of the fresh

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spells in the 2004 panel start on the first month of the reference period. For the survival analysis in (5.1) long term trends (e.g. over a year) are the focus, seam bias will not be of great concern. However, it will be taken into consideration when interpreting results in subsection (5.2) where variations at the monthly level are of interest.

Second, Medicaid and CHIP coverages cannot be reliably distinguished.[12] Therefore, I will use public insurance coverage which encapsulates both, and the phrase "public insurance" will be used interchangeably with Medicaid/CHIP or simply Medicaid. When computing the value of the benefit notch, I will use the CHIP government spending per enrollee, which is lower than that of Medicaid. The implication of ignoring the higher benefit notch Medicaid applicants face will be discussed in section (6).

Third, identifiers of families in less populated states are missing from the 2001 panel. Families in Maine and Vermont share the same state identifier as well as those in North Dakota, South Dakota and Wyoming. Because these states have different Medicaid/CHIP policy parameters, they are excluded from analyses. Due to the larger sample size of the 2004 panel, all fifty states plus the District of Columbia have their own identifier, and therefore all states can be included.

The main analysis sample consists of children who started a public insurance spell during the SIPP panel. The restriction to those with "fresh" spells (as opposed to the left-truncated spells that start with the child's first appearance in the panel) comes from the necessity of identifying when the family applied for benefit, which is not possible with the left-truncated spells. In addition, children younger than the age of 1 and children whose families moved to another state during the spell are excluded from the analysis sample. Infants are excluded because most states have been extending presumptive eligibility to infants since the 1990's and children whose families moved across the states pose a challenge in assigning Medicaid/CHIP parameters. As shown in Table (2), the analysis sample consists of–for the 2001 and 2004 panels respectively–7158 and 8321 fresh spells in total and 3096 and 3312 fresh spells from children in the states that offer continuous eligibility.

Nuclear families for each child are constructed using information on the relationship to household and family reference person (head). In cases where a child and his or her parent(s) live with other adults, however, families include only the children and parent(s) of the appropriate subfamily. This definition

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[12]Amy Steinweg at the U.S. Census Bureau noted in a correspondence that "respondents rarely know with certainty whether their child is in Medicaid or CHIP... We found this out with the 2004 SIPP instrument, where question order happened to be revised so that CHIP was asked about before Medicaid. Here we observed that respondents were most likely to answer the question asked first, resulting in higher reported levels of CHIP than of Medicaid for Panel 2004)".

corresponds to the family assistance unit that would be potentially eligible for Medicaid/CHIP. Variables such as family income are then calculated by aggregating over individual family members.

The state level Medicaid/CHIP data are extracted from reports issued and database maintained by various organizations. The policy parameters (e.g. continuous eligibility, presumptive eligibility, income eligibility cutoffs, etc) come from NGA (2000-2008), Kaiser (2002-2011) and CMS (Various Years). Medicaid/CHIP spending and enrollment data come from the Kaiser Foundation State Health Facts database and the CMS Medicaid Statistical Information Statistics System.

Table XXX presents summary statistics.

5   Empirical Results

5.1   Public Insurance Spells

An intuitive prediction from the labor supply models' above is that children in states that offer 12 months of continuous eligibility are expected to remain covered for the entire 12-month duration. In this section, I examine the duration of children's public insurance spells using SIPP data. The descriptive evidence suggests that only about half of reported coverages lasted more than 12 months even in states that offer 12-month continuous eligibility.

Figure (1) plots the Kaplan-Meier survival function for all Medicaid/CHIP spells by whether or not children resided in a state offering 12 months of continuous eligibility for the 2001 and 2004 panels respectively. Two patterns deserve attention when interpreting the results. First, seam bias is conspicuous at the months that are multiples of 4. As mentioned in section (4), most of the fresh spells start on the first month in a reference period, and analogously most of the spells either end or are censored on the fourth month in a reference period. Consequently, most of the observed spell lengths are multiples of four. Although the presence of the seam bias confounds the duration analysis for short term spells, it has only minor effect on the estimate of the probability of maintaining coverage at month 12. It does not change the fact that about half of the fresh public insurance spells (55% for the 2001 panel and 51% for the 2004 panel) do not last as long as 12 months.

Second, there is little difference in the spell durations between children in states that offer continuous eligibility and those that do not. Various statistical tests show that the difference between the two survival functions is marginal at the 5% level. For the 2001 panel, the p-values associated with the log-rank,

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Wilcoxon, Tarone and Ware and Peto tests are 0.11, 0.03, 0.05 and 0.06, respectively. For the 2004 panel, the p-values are 0.06, 0.02, 0.05 and 0.04, respectively. One reason the two survival functions are so similar is that the underlying recertification period for many of the states that do not offer continuous eligibility are 12 months long as well Kaiser (2002-2011). Even though change reporting may be required for families in these states when their circumstances change, its enforcement may be weak. Another reason is that the similarity may be partially driven by states in each group that offers presumptive eligibility to children. It is reasonable to suspect that there are more short spells in the presumptive eligibility states because of the children that obtain coverage temporarily but later found to be ineligible. Dropping the states that offer presumptive eligibility shows that the survival function for children in the 12-month continuous eligibility states lie slightly above those that are not. Stratified equality tests show that within the states that offer presumptive eligibility, the survival function are very similar between states that offer continuous eligibility and those that do not (p-value larger than 0.3 for the 2001 panel and larger than 0.65 for the 2004 panel). The survival functions are again marginally different in states that do not offer presumptive eligibility with p-values ranging from 0.02 to 0.05 for the four statistical tests in both the 2001 and 2004 panel, yet the difference is not economically significant as the survival functions are far from that showing 12 months of continuous coverage for most children.

Just as there were short coverages, there were short gaps before the start of a spell as well. Short coverage gaps are not inconsistent with what is found using administrative data. For example, using administrative data from Ohio, Fairbrother et al. (2011) 40% of the children whose coverage was not renewed at month 12 re-enroll within a short period. However, it does cause concern regarding the reliability of identifying the start of a spell. Although it is rare for families to report coverage while they are not on public insurance (Card et al. (2004)), underreporting coverage at a particular month during a long spell will lead to the false identification of starting a fresh spell.

To address this potential measurement error problem, I construct a subsample consisting of children who were reported not to be covered by public insurance for 12 consecutive months before the spell. This subsample will be henceforth titled the "long gap" sample. As shown in Table (2), the 2001 panel long gap sample includes 501 spells from 501 children in 302 sample units and the 2004 panel long gap sample includes 815 spells from 804 children in 494 sample units. Note that the long gap sample contains mostly of single spells by virtue of its construction. Figure (2) plots the survival function in the long gap sample–the general pattern of Figure (1) remains and the survival functions between the two group of states–those with

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and without continuous eligibility–are not statistically different.

I have examine whether the following candidates can explain the children's public insurance coverage being prematurely dropped: parents' Medicaid coverage dropped; TANF coverage dropped; acquired private insurance. As it turns out, these factors cannot account for short public insurance spells of children–only about 2 percentage points of the children in the sample switch off TANF when going off Medicaid, 10-20 percentage points of children switch off public insurance with their mother and only about 10 percentage points of the Medicaid/CHIP leavers acquire private insurance when switching off public insurance. It is possible that CHIP enrollees may choose not to pay the monthly premium and hence discontinue coverage.

To summarize, the probability that a fresh public insurance spell duration exceeds 12 months is approximately 50 percent in all samples. This is consistent with the disenrollment estimates reported in the literature (e.g. Shulman et al. (2006)). There appear to be short coverage spells as well as short gaps between spells. Theses transitions will be taken into account in the next subsection when I describe families' income movement around the start of a child's public insurance spell.

5.2   Family Income Responses

In the subsection, I present descriptive evidence on families' income response over their childrens' Medicaid/CHIP spell. I follow a flexible specification adopted by Jacobson et al. (1993). Specifically I estimate

$$y_{it} = \omega_i + \gamma_t + \sum_{|k| \geqslant m} D_{it}^k \delta_k + \varepsilon_{it} \tag{11}$$

where ω_i and γ_t are individual and calendar month fixed effect respectively and D_{it}^k is a set of dummy variables indicating months after the start of a public insurance spell. $D_{it}^k = 1$ if child i started her public insurance spell at month $t - k + 1$. As a special case, a child with $D_{it}^0 = 1$ started her public insurance spell in month $t + 1$, and another child with $D_{it}^1 = 1$ started her spell in month t. Month 0 is the omitted category and the δ_k's measure the difference in the average outcome k months after the start of a spell relative to the value at the beginning of the spell. I examine the income and labor supply responses 24 months before and after the beginning of a spell. Since families were interviewed for 36 months in the 2001 panel and 48 months in the 2004 panel, it is possible to allow $m = 35$ for the 2001 panel and $m = 47$ for the 2004 panel. However, there are few families who start a spell at the second month or the last month of the panel which render the estimation of δ_m and δ_{-m} imprecise for large m. Moreover, choosing $m = 24$ is sufficient for the purpose of

testing the labor supply models above as it covers two years of data surrounding an additional recertification months ($m = 12$).

According to the theoretical models in Section (3), the δ_k's are expected to be positive for k not being a multiple of 12 (eligibility check points) if there is strategic behavior. However, even in the absence of any strategic behavior, one might expect a mechanical dip as pointed out by Ashenfelter (1978) and Ashenfelter and Card (1985). For example, if selecting into public insurance is based on $y_{i(t_0-k)} < \bar{y}$ where t_0 is the month of starting the spell and that the ε_{it}'s are serially correlated, then a dip and rebound may be expected based purely on mean reversion, though not as abrupt as predicted by the strategic behavior. Given the existence of the Ashenfelter dip mechanism, the dip and rebound due to the strategic behavior should be more pronounced.

As noted in subsection (5.1), many families in states that provide 12-months of continuous eligibility do not report coverage for all 12 months after beginning the spell. Likewise, many families only experience short gaps leading up to the beginning of a public insurance spell. This makes it difficult to interpret the δ_k's in an event-study framework, which usually calls for single status transitions. Therefore, I will truncate the analysis sample for each child and focus on the single transitions. Specifically, let k^+ denote the first month after month 0 the child switches off public insurance, and let k^- denote the last month before month 0 the child was covered by public insurance. I will discard all observations after k^+ and those before k^-.

Figure (3) and Figure (4) plot the movement of (unweighted) average family income over the 24 periods around the beginning of a public insurance spell. Both point estimates and standard errors are shown where the standard errors are clustered at the sample unit level. None of the figures show a pronounced dip-and-rebound in the six months before and after the spell start. For the 2001 Panel, the income trend leading up to the beginning of spell is flat in both samples, however, the 95% confidence interval does not rule out a downward trend. For the full sample specification, the income increases gradually especially after 12 months, but the period immediately following the spell start shows no rebound. Even the upward income trend between 12 months and 24 months after the spell is not statistically different from zero, and it disappears altogether when restricting to the long gap sample in Figure (4). In the 2004 panel, the income process shows a persistent downward trend in both samples, and the estimates of δ_k are significantly positive for many $k < 0$ and negative for many $k > 0$.

Unfortunately, SIPP does not collect information on hours worked on a monthly level but usual hours worked are reported for the entire wave. If the labor supply models presented above are true and imply only

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supply movements immediately before and after a public insurance spell, then using the usual hours worked variable may cloud the dip-and-rebound patter. Instead, I describe the movement of the less refined labor supply variables at the monthly level before and after the start of a public insurance spell. In particular, I construct a dummy variable indicating whether or not the head of the family–defined to be father in a two-parent family and all single-parent family heads–worked more than 35 hours for all weeks during a month, and plot its movement in Figure (5) and Figure (4) for the full sample and long gap sample respectively. In all the figures, there is a slight downward trend for the two years before the spell start, and it continues in the 2001 panel for at least 6 months in both samples. The trend flattens after the insurance spell in the full sample of the 2004 panel, but the downward trend continues as with the 2001 panel when I restrict to the long gap sample.

In summary, both the labor supply theory and mechanical mean reversion predict a dip-and-rebound in family income and labor supply around the start of a public insurance spell. However, I do not find such patterns using the various subsamples of the SIPP 2001 and 2004 panel. In some cases, the average income are significantly below that of month 0, but small sample size lead to a 95% confidence interval whose upper bound is positive. Therefore, strategic behavior cannot be strictly ruled out. In the next section, I present results from calibrating the labor supply models using various elasticity measures and gauge whether the upper bounds of the 95% confidence interval is consistent with the quantitative theoretical predictions.

6 Calibration

In this section, I provide quantitative predictions by calibrating the simple models presented in section (3). Specifically, I focus on the dynamic model where the labor supply choice is continuous and the flow utility is of the form (3.1) and (7) and attempt to find the labor supply elasticity parameter that is consistent with the empirical evidence. The case where labor supply choices are discrete is being currently investigated.

For the calibration exercise, I generate 100,000 families for whom I assign the taste parameter n, income eligibility cutoff γ, size of the benefit notch as well as the tax rate t. The distribution of n is calibrated using family income distribution from SIPP data. In particular, recall that the optimal pretax income choice for a family with quasi-linear utility and taste parameter n facing the budget constraint $C = (1-t)Z$ is $Z^* = (1-t)^e n$. For each family in SIPP, I assign the marginal federal income tax rate based on their income

and family composition.[13] Using families with children residing in states that offer continuous eligibility, I impute their value of n as $Z^*/(1-t)^e$ where Z^* is family income. Assuming log-normality, I estimate the mean and standard deviation of the distribution of $\log(n)$ for each value of e. For the plausible range of e values, the distribution of n changes only slightly with a mean around 8.2 and a standard deviation of 1.2 for both the SIPP 2001 and 2004 panels. In the case where the flow utility is of the form (7), $n = \frac{Z^*}{(1-t)^e}[(1-t)Z^*]^{\rho e}$, but its distribution can be calibrated accordingly (note that if the distribution of Z^* is approximately log-normal, then so is that of n).

In this section, I consider only the CHIP benefit notch, which is the highest benefit notch a family with children faces. I will discuss below the implication of ignoring other notches in the lower part of the budget constraint. An estimate of the benefit notch value g comes from the CHIP spending data collected by the Kaiser Foundation and the Center for Medicare and Medicaid Services. The spending variable excludes beneficiary and third-party payment and should reflect expected government subsidy. Unfortunately, state-by-state spending per child enrollee figures are not readily available from either source for years earlier than 2004. Therefore, I use the average per-enrollee-spending for the entire U.S. as a measure of the notch, which had increased from \$835 in 2001 to \$1217 in 2007 in nominal terms (approximately 5% annual growth rate). The monthly benefit amount per child g is $\frac{1}{12}$ of the annual spending per-enrollee spending, and the size of the notch a family faces is the g times the number of children they have. The marginal distribution of the number of children in each family in the simulation sample mimics that in SIPP–approximately 32%, 37% and 21% of families have 1, 2 and 3 children in both the 2001 and 2004 panels of SIPP. Finally, the CHIP income eligibility cutoff γ is extracted from the cutoffs families face in continuous eligibility states in SIPP sample.

I show calibrations of two models–one with no income effects (the quasi-linear case) and the other with $\rho = 1$. As Chetty (2006) summarizes, the average value of ρ from empirical literature is 0.74 and in general less than 1 for micro studies, which is the basis of choosing $\rho = 1$. As expected, when income effects are taken into account, the magnitudes of the dip and the rebound are smaller as the increase/decrease in income render agents demand more/less leisure. Also, the rebound is smaller in magnitude than the dip, reflecting the demand for more leisure as children in families acquire public insurance.

Results from the SIPP 2001 and 2004 panels are presented in Table (4) and Table (5) respectively. In the

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[13]Specifically, I assume that parents in a dual-headed family file jointly and claim the deduction accordingly and that all families claim standard deductions.

2001 panel truncated full sample, elasticities larger than or equal to 0.05 can be effectively ruled out by both the point estimates and the upper 95% confidence interval. In the long gap sample, however, while the point estimates are mostly negative and rule out all elasticities larger than 0.05, the upper 95% confidence interval for the month 4 estimate may be consistent with elasticities larger than 0.2 in the no income effect case and those larger than 0.5 in the income effect case. In the 2004 panel, the point estimates and the upper 95% confidence interval for months after spell start in both the full sample and the long gap sample rule out an elasticity of 0.05 and greater in both models. In general, the empirical evidence points to small elasticities.

Note that the theoretical prediction of dip and rebound magnitudes may even be an underestimate for two reasons. First, as pointed out in section (5.2), the existence of the Ashenfelter dip should accentuate the dip and rebound. Second, because of the high income cutoff of CHIP, there are many programs agents eligible for CHIP may qualify for if they reduce their income. For example, even though a five-year-old child in a family with income at 140% of the FPL is eligible for CHIP in practically all states, the family will face a more generous transfer by reducing their income to 133% of the FPL. In this case, the child will qualify for Medicaid in every state and the family will incur no premium payments and the lowest co-pay. If the family is willing to reduce their income further to 130% of the FPL, they will also gain eligibility for the SNAP program (formerly Food Stamp).

The implied labor supply elasticities are small and consistent with those estimated from micro data using nonlinearities in the budget constraint (e.g. Saez (2010) and Chetty et al. (2011)). The findings likely imply the friction involved in adjusting income or equivalently the small perceived value of CHIP to those families above the eligibility cutoff. I will explore alternative modeling strategies in future work.

7 Optimal Continuous Eligibility Period–A Discussion

In this section, I consider the implication of the empirical findings in Section (5) on the optimal continuous eligibility period. Rather than allowing the choice set of continuous eligibility period τ to be any integer value, I restrict the candidates to be $\tau = r$ and $\tau = 2r$. In reality, all states provided continuous eligibility either for 6 or 12 months. So the policy question I attempt to address is: given what is observed from the states that allow 12 months of continuous eligibility, should they reduce it to 6 months or should the 6-month states increase it to 12 months. r, therefore, is equal to 6 in the policy context. I will show below that the extent of the strategic behavior is informative of the design problem and that the 12-month recertification

period dominates that of 6 months as suggested by the point estimates from the empirical evidence.

The government solves the problem

$$\max_{t,\tau} \int \Psi(U(n))f(n)dn$$

where Ψ is a weighting function that reflects the preference of the policy maker.

$$U(n) \quad = \quad \max_{\{C_s,Z_s\}} \sum_{s=0}^{T-1} \beta^s u(C_s, Z_s; n)$$

with C_s and Z_s being the optimal pre-tax and post-tax incomes respectively. Here, $U(n)$ is the maximum utility when resulting from the dynamic problem of T periods laid out in (3.2) with the flow utility being (1) except agents choose $Z = wH$ directly. Also, I allow the transaction cost of filing application ϕ for the agents when eligibility is checked as per Currie et al. (2001). As seen from section (3.2), agents will participate in the program if and only if $n < \bar{n}(\tau)$ where $\bar{n}(\tau)$ is the highest type of agent participating in the program. In effect, agent of type $\bar{n}(\tau)$ is indifferent between bunching at the notch of size $\tau g - \phi$ (assuming $\beta \approx 1$ from here on) and choosing her desired pre-tax income in the absence of the notch.

Therefore, the government's budget constraint is (assuming interest rate is 0)

$$\int \sum_{s=0}^{T-1} t Z_s(n)f(n)dn = (Tg + m\frac{T}{\tau})\Pr(n < \bar{n}(\tau))$$

where f is distribution of agent type n, m is the monitoring cost of verifying eligibility each time.[14] Without loss of generality, let there be $T = 2r$ periods.

Qualitatively, increasing τ from r to $2r$ ceteris paribus has the following effects:

1. $\Pr(n < \bar{n}(\tau))$ increases–more agents will participate in the program.

2. $m\frac{T}{\tau}$ decreases from $2m$ to m–monitoring cost per person is cut in half.

3. Tax revenue gain from agents with $n \in (\bar{n}_\gamma, \bar{n}(\tau))$–their labor supply decision is no longer distorted in period $s = r$.

4. Tax revenue loss from agents with $n \in (\bar{n}(\tau), \bar{n}(2\tau))$–they will bunch at γ at $s = 0$ to gain program eligibility.

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[14]For an arbitrary T, the monitoring cost should be $m\lfloor\frac{T}{\tau}\rfloor$. The greatest integer operator is ignored because I will choose T to be a multiple of τ later on.

5. Reduced fixed cost of application from 2ϕ to ϕ−utility gains for those with $n \in (0, \bar{n}(2r))$.

Consider the net budgetary effect of lengthening τ from r to $2r$. The net expenditure increase in public insurance benefit paid is

$$(Tg)[\Pr(n < \bar{n}(2r)) - \Pr(n < \bar{n}(r))] \tag{12}$$

the net expenditure reduction in monitoring cost is

$$2m\Pr(n < \bar{n}(r)) - m\Pr(n < \bar{n}(2r)) \tag{13}$$

and the net tax revenue loss is

$$\int_{n\gamma}^{\bar{n}(2r)} t(Z^0(n) - \gamma)f(n)dn - \int_{n\gamma}^{\bar{n}(r)} t(Z^0(n) - \gamma)f(n)dn = \int_{\bar{n}(r)}^{\bar{n}(2r)} t(Z^0(n) - \gamma)f(n)dn \tag{14}$$

where $Z^0(n) = n(1-t)^e$ is the income an agent of type n would have chosen in the absence of the notch.

(12) is equal to

$$Tg\int_{\bar{n}(r)}^{\bar{n}(2r)} f(n)dn \approx Tg(\bar{n}(2r) - \bar{n}(r))\frac{f(\bar{n}(2r)) + f(\bar{n}(r))}{2}$$

where the approximation is by Trapezoid Rule. When $e \approx 0$, the agents' indifference curves are close to Leontief. Consequently $\bar{n}(2r) \approx \frac{2rg - \phi}{1-t} + \gamma$ and $\bar{n}(r) \approx \frac{rg - \phi}{1-t} + \gamma$ where $2rg - \phi$ and $rg - \phi$ are the notch sizes when $\tau = 2r$ and $\tau = r$ respectively. In order to simplify the analyses, I will assume from now on that $\phi \approx 0$ for the agents with $n \geqslant \gamma$, namely the fixed cost of applying for benefits is low for high income agents. It follows that (12) is approximately

$$\frac{(2rg)(rg)}{1-t}\frac{f(\bar{n}(2r)) + f(\bar{n}(r))}{2} \tag{15}$$

Similarly, (14) is approximately

$$\frac{t[(rg)(2rg)f(\bar{n}(2r)) + (rg)(rg)f(\bar{n}(r))]}{2(1-t)^2} \tag{16}$$

I will assume that the distribution of n is such that increasing τ to $2r$ from r brings a reduction in monitoring cost, i.e. (13)>0. Therefore, (12)+(14) represents the maximum deficit associated with doubling the recertification period.

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Now I will show that (12)+(14) is bounded above by the average income dip provided a weak sufficient condition on the distribution of n is satisfied. The income dip observed empirically is

$$\int_{n_\gamma}^{\bar{n}(2r)} (Z(n) - \gamma) f_{n|n<\bar{n}(2r)}(n') dn' = \frac{1}{\Pr(n < \bar{n}(2r))} \int_{n_\gamma}^{\bar{n}(2r)} (Z(n) - \gamma) f(n) dn \tag{17}$$

where the conditional distribution $f_{n|n<\bar{n}(2r)}$ is used because the analysis sample is restricted to families that participated in Medicaid/CHIP. Approximating (17) by Trapezoid Rule gives

$$\frac{(2rg)(rg) f(\bar{n}(2r))}{(1-t)^2 \Pr(n < \bar{n}(2r))}$$

After algebra arrangement, (17)-[(12)+(14)] is approximately

$$\frac{(2rg)(rg)[(2-p) f(\bar{n}(2r)) - (1 - \frac{1}{2}t) p f(\bar{n}(r))]}{2(1-t)^2 p}$$

where $p = \Pr(n < \bar{n}(2r))$. If $\frac{f(\bar{n}(2r))}{f(\bar{n}(r))} \geqslant \frac{(1 - \frac{1}{2}t)p}{2-p}$, (17)-[(12)+(14)]$\geqslant 0$. The condition $\frac{f(\bar{n}(2r))}{f(\bar{n}(r))} \geqslant \frac{(1 - \frac{1}{2}t)p}{2-p}$ is satisfied whenever $f(\bar{n}(2r)) \geqslant f(\bar{n}(r))$ (for example f is distributed uniformly or when $\bar{n}(2r)$ and $\bar{n}(r)$ are both located on the increasing range of f). It is also satisfied under the particular parameter values in (6), which are based on policy variables and income distributions in SIPP.

If no significant dip is observed, then the budgetary pressure created by increasing the continuous eligibility period is small. Provided that there are savings from monitor cost reduction, the government's budget constraint is still satisfied. Since setting $\tau = 2r$ will lower the fixed application cost for the agents with low n, $\tau = 2r$ dominates $\tau = r$. In the policy context, those states still with 6 months of renewal period, namely Georgia and Texas should consider halving the renewal frequency, and those currently offering 12 months of continuous eligibility should not switch back to six months as in the case of–for example–Connecticut, Indiana, Nebraska, Washington and New Mexico in the early 2000's (Ross and Cox (2003), Ross and Cox (2004), Heberlein et al. (2011)).

8   Conclusion

This paper considers the impact of the continuous eligibility provision in income-tested programs on the labor supply responses of program participants. In particular, neo-classic labor supply models predict that

the provision provides dynamic opt-in incentives wherein families lower their income to gain program eligibility, acquire government-provided benefits for the continuous eligibility period and revert back to their "optimal" consumption bundle. Using 2001 and 2004 panels from SIPP, I follow a flexible event-study specification and find no evidence of the dip-and-rebound strategic behavior in family income around the time a child initially gains coverage of public insurance.

I calibrate the neo-classic dynamic labor supply model both with and without income effects by using family income and composition information from SIPP, Medicaid/CHIP policy parameters and income tax rates. Comparing the magnitudes of the predicted strategic behavior to those observed empirically points to small labor supply elasticities. In many subsamples, elasticities higher than 0.05 can be effectively ruled out even when using the upper bound of the 95% confidence interval of the empirical point estimates.

The lack of evidence supporting strategic behavior suggests that the labor supply distortion created by the continuous eligibility provision is small. Intuitively, this means that the 12-month continuous eligibility period is not longer than optimal. As I show formally in an optimal design framework, it may be beneficial for states currently allowing a 6-month renewal period to extend it to 12 months as it may lower the renewal cost for the program participants and government without inducing many high income families to opt in.

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Appendix

States Providing Continuous Eligibility and Presumptive Eligibility

Continuous Eligibility

- The states that provided 12 months of continuous coverage in the sample period (2001-2007) are: Alabama, California, Washington D.C., Idaho, Illinois, Iowa, Kansas, Louisiana, Mississippi, Maine, Michigan (after Jan 2003), New York, North Carolina, Ohio, Pennsylvania, South Carolina, Washington, and West Virginia (after Oct 2002).

- The states that did not provide 12 months of continuous eligibility are: Colorado, Georgia, Hawaii, Kentucky, Missouri, Montana, Nevada, New Hampshire, North Dakota, Oklahoma, Rhode Island, South Dakota, Tennessee, Texas, Utah, Vermont, Virginia, Wisconsin, and Wyoming.[15]

- In the states that remain–Alaska, Arizona, Arkansas, Connecticut, Delaware, Florida, Indiana, Maryland, Massachusetts, Minnesota, Nebraska, New Jersey, New Mexico, and Oregon–the rules for continuous eligibility are complicated with changes or different implementations for different programs during my sample period, and are thus dropped from my analysis sample.

Presumptive Eligibility

- The states that provided presumptive eligibility are California, Connecticut, Florida, Illinois, Massachusetts, Michigan, Missouri, Nebraska, New Hampshire, New Jersey, New York and Oklahoma.

Discrete Labor Supply Choices

In this section, I show that there is no income bunching but a discontinuity in the income density at the eligibility cutoff in the case where finitely many choices of hours are allowed. Formally, let the menu of labor supply choices be $\{h_1, h_2, ..., h_d\}$ where $0 = h_1 < h_2 < ... < h_d = 1$. For each w, let $\bar{n}_{h_i,h_{i+1}}(w)$, $i = 1, 2, ...d$, be the type of agent indifferent between choosing $H = h_i$ and $H = h_{i+1}$, and it follows that agents facing wage w of type $n \in (\bar{n}_{h_{i-1},h_i}(w), \bar{n}_{h_i,h_{i+1}}(w)]$ choose $H = h_i$.[16] $\bar{n}_{h_i,h_{i+1}}(w)$ varies smoothly with w except when $wh_i = \gamma$ and $wh_{i+1} = \gamma$. There is a discontinuous increase when $w = \gamma/h_{i+1}$ and a discontinuous decrease

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[15]Note that many of the states that did not provide continuous eligibility allowed a 12-month renewal period.

[16]$\bar{n}_{h_d,h_{d+1}}$ is defined to be ∞.

when $w = \gamma/h_i$. The c.d.f. of Z at $z > 0$ is

$$F_Z(Z < z) = \Pr(H = 0) + \sum_{i=2}^{d} \int_0^{z/h_i} \int_{\bar{n}_{h_{i-1},h_i}(w')}^{\bar{n}_{h_i,h_{i+1}}(w')} f_{n,w}(n',w')dn'dw'$$

F_Z is continuous at γ because $\bar{n}_{h_i,h_{i+1}}(w)$ is right continuous for all i and that $f_{n,w}$ is continuous. However,

$$\lim_{z\uparrow\gamma} f_Z(z) = \sum_{i=2}^{d} \frac{1}{h_i} \int_{\bar{n}^l_{h_{i-1},h_i}(\gamma/h_i)}^{\bar{n}^l_{h_i,h_{i+1}}(\gamma/h_i)} f_{n,w}(n', \frac{\gamma}{h_i})dn'$$

$$\lim_{z\downarrow\gamma} f_Z(z) = \sum_{i=2}^{d} \frac{1}{h_i} \int_{\bar{n}^r_{h_{i-1},h_i}(\gamma/h_i)}^{\bar{n}^r_{h_i,h_{i+1}}(\gamma/h_i)} f_{n,w}(n', \frac{\gamma}{h_i})dn'$$

where

$$\bar{n}^l_{h_i,h_{i+1}}(\frac{\gamma}{h_i}) \equiv \lim_{w\uparrow(\frac{\gamma}{h_i})} \bar{n}_{h_i,h_{i+1}}(w) > \lim_{w\downarrow(\frac{\gamma}{h_i})} \bar{n}_{h_i,h_{i+1}}(w) \equiv \bar{n}^r_{h_i,h_{i+1}}(\frac{\gamma}{h_i})$$

$$\bar{n}^l_{h_{i-1},h_i}(\frac{\gamma}{h_i}) \equiv \lim_{w\uparrow(\frac{\gamma}{h_i})} \bar{n}_{h_{i-1},h_i}(w) < \lim_{w\downarrow(\frac{\gamma}{h_i})} \bar{n}_{h_{i-1},h_i}(w) \equiv \bar{n}^r_{h_{i-1},h_i}(\frac{\gamma}{h_i})$$

It follows that

$$\int_{\bar{n}^r_{h_{i-1},h_i}(\gamma/h_i)}^{\bar{n}^r_{h_i,h_{i+1}}(\gamma/h_i)} f_{n,w}(n', \frac{\gamma}{h_i})dn' < \frac{1}{h_i} \int_{\bar{n}^l_{h_{i-1},h_i}(\gamma/h_i)}^{\bar{n}^l_{h_i,h_{i+1}}(\gamma/h_i)} f_{n,w}(n', \frac{\gamma}{h_i})dn'$$

for all i, and therefore $\lim_{z\uparrow\gamma} f_Z(z) > \lim_{z\downarrow\gamma} f_Z(z)$.

Table 1: Total Counts of Individuals and Sample Units and Those Covered by Public Insurance During Panel

(a) Total Individual and Sample Unit Counts

	No. of Individuals		No. of SU's	
	2001	2004	2001	2004
Total	104053	131549	35106	43540
Children Living with Parent(s)	29549	37333	14323	17780

(b) Individual and Sample Unit Counts: on Public Insurance During Panel

	No. of Individuals on Public Insurance		No. of SU's on Public Insurance	
	2001	2004	2001	2004
Total	23152	33986	10001	14454
Children Living with Parent(s)	11109	17458	5272	8235

Notes: Panel (a) shows the total number of children, individuals and sample units in the 2001 and 2004 SIPP panels. The bottom row of the right column shows number of sample units that have children living with parents. Panel (b) shows the total number of children, individuals and sample units ever on public insurance. The top row of the right column shows the number of sample units that had at least a member on public insurance during the panel, and the bottom row shows the number of sample units that had at least a child on public insurance during the panel.

Table 2: Public Insurance Spell, Child and Sample Unit Counts by Spell Types, Continuous Eligibility States and in Analysis Sample

(a) Public Insurance Spell, Child and Sample Unit Counts by Spell Types and Continuous Eligibility States

	Total No. of Public Insurance Spells		No. of Kids with Public Insurance Spells		No. of SU with Kids on Public Insurance	
	2001	2004	2001	2004	2001	2004
Pub Insurance spells	16109	23109	10656	17190	5024	8149
Left-Truncated Spells	8402	13996	7407	11929	3683	5848
Fresh Spells	7707	9113	5759	7850	3033	4390
Fresh Spells Ex. Infants & State Movers	7158	8321	5360	7188	2866	4099
12-Month Continuous Elig.	3096	3312	2310	2843	1257	1642
No Continuous Elig.	2270	2535	1742	2217	934	1255
Other States	1792	2474	1319	2147	685	1223

(b) Public Insurance Spell, Child and Sample Unit Counts in Analysis Sample

	Subsample Public Insurance Spells		No. of Kids with Public Insurance Spells		No. of SU with Kids on Public Insurance	
	2001	2004	2001	2004	2001	2004
Fresh Spells: States with 12 Months of Continuous Elig.	3096	3312	2310	2843	1257	1642
Long Gap Sample	501	815	501	804	302	494
Long Gap and No Presumptive Elig.	232	389	232	382	134	232

Notes: This table shows the number of observations excluded in each subsample. The full analysis sample consists of 3096 spells from 2310 children in 1257 sample units in the 2001 panel and 3312 spells from 2843 children in 1642 sample units in the 2004 panel. For the 2001 panel, the analysis sample is reached by excluding 8402 left-truncated Spells, 549 spells from infants and children in families that moved during the panel, and 4062 spells from children not in the states that provide continuous eligibility. For the 2004 panel, the analysis sample is reached by excluding 13996 left-truncated Spells, 792 spells from infants and children in families that moved during the panel, and 5009 spells from children not in the states that provide continuous eligibility.

Table 3: Variable Means for Children in the Analysis Sample

	2001 Panel				2004 Panel			
	Full Sample		Long Gap Sample		Full Sample		Long Gap Sample	
	Month 0	Month 1	Month 0	Month 1	Month 0	Month 1	Month 0	Month 1
Age	8.17	8.26	9.15	9.14	8.76	8.85	9.27	9.36
Female	0.5	0.5	0.5	0.5	0.48	0.48	0.5	0.5
Black	0.25	0.25	0.2	0.2	0.17	0.17	0.12	0.11
Family Size	4.1	4.1	4.12	4.13	4	4	4.2	4.2
Single Parent Family	0.49	0.49	0.38	0.38	0.41	0.41	0.27	0.27
Family Income (nominal)	2139	2048	2914	2847	3585	3615	4860	4828
Family Income (in 2010 $)	2582	2468	3481	3396	3990	4014	5312	5268
Income Eligibility Cutoff	3359	3372	3403	3415	3684	3697	3919	3930
Fraction without Earned	0.11	0.12	0.05	0.07	0.08	0.07	0.04	0.05
On Welfare	0.05	0.07	0.01	0.02	0.03	0.04	0.01	0.01
Covered by SSI	0.01	0.02	0.004	0.007	0.01	0.02	0.01	0.01
On Medicaid	0	1	0	1	0	1	0	1
On Private Insurance	0.35	0.26	0.42	0.31	0.5	0.46	0.63	0.57
Mom on Medicaid	0.19	0.36	0.07	0.24	0.18	0.27	0.01	0.15
Dad on Medicaid	0.08	0.16	0.03	0.1	0.07	0.12	0.04	0.07
Mom on Private Insurance	0.36	0.3	0.47	0.41	0.54	0.52	0.69	0.68
Dad on Private Insurance	0.43	0.38	0.52	0.47	0.66	0.65	0.77	0.77
Mom on UI	0.02	0.02	0.03	0.04	0.01	0.02	0.02	0.02
Dad on UI	0.05	0.05	0.07	0.08	0.03	0.02	0.02	0.03

Notes: Variable means for children in the analysis sample right before and during the first month of public insurance spell.

Table 4: Empirical *vs.* Predicted Income Dip and Rebound for Various Labor Supply Elasticities: SIPP 2001 Panel

	Empirical Measures of Income				Simulated Measures of Income					
	Truncated Full Sample		Truncated Long Gap Sample		No Income Effect (ρ=0)			Income Effect (ρ=1)		
Month in Spell	Point Estimates	Upper 95% CI	Point Estimates	Upper 95% CI	e=0.5	e=0.2	e=0.05	e=0.5	e=0.2	e=0.05
-5	32	181	17	554	814	487	329	450	331	231
-4	11	43	14	458	814	487	329	450	331	231
-3	-56	19	-195	251	814	487	329	450	331	231
-2	-55	49	-169	177	814	487	329	450	331	231
-1	-22	35	-110	239	814	487	329	450	331	231
0	0	0	0	0	0	0	0	0	0	0
1	-62	70	-199	168	814	487	329	391	301	223
2	-36	102	-157	236	814	487	329	391	301	223
3	-15	130	-118	273	814	487	329	391	301	223
4	-3	130	-245	509	814	487	329	391	301	223
5	-46	170	-197	149	814	487	329	391	301	223
6	-20	220	-440	172	814	487	329	391	301	223

Notes: Model calibration assuming continuous labor supply choice. Empirical estimates are based on specification (11) where the point estimates trace out average family income around the start of a child's public insurance spell. Model parameters are calibrated using 2001 SIPP panel income data, federal income tax rates, published Medicaid/CHIP eligibility cutoffs and CHIP spending data. Note that the implied Hicksian labor supply elasticity when $\rho = 1$ is $\frac{e}{\rho e+1}$, which is 0.33, 0.16 and 0.048 for e =0.5, 0.2 and 0.05 respectively.

Table 5: Empirical vs. Predicted Income Dip and Rebound for Various Labor Supply Elasticities: SIPP 2004 Panel

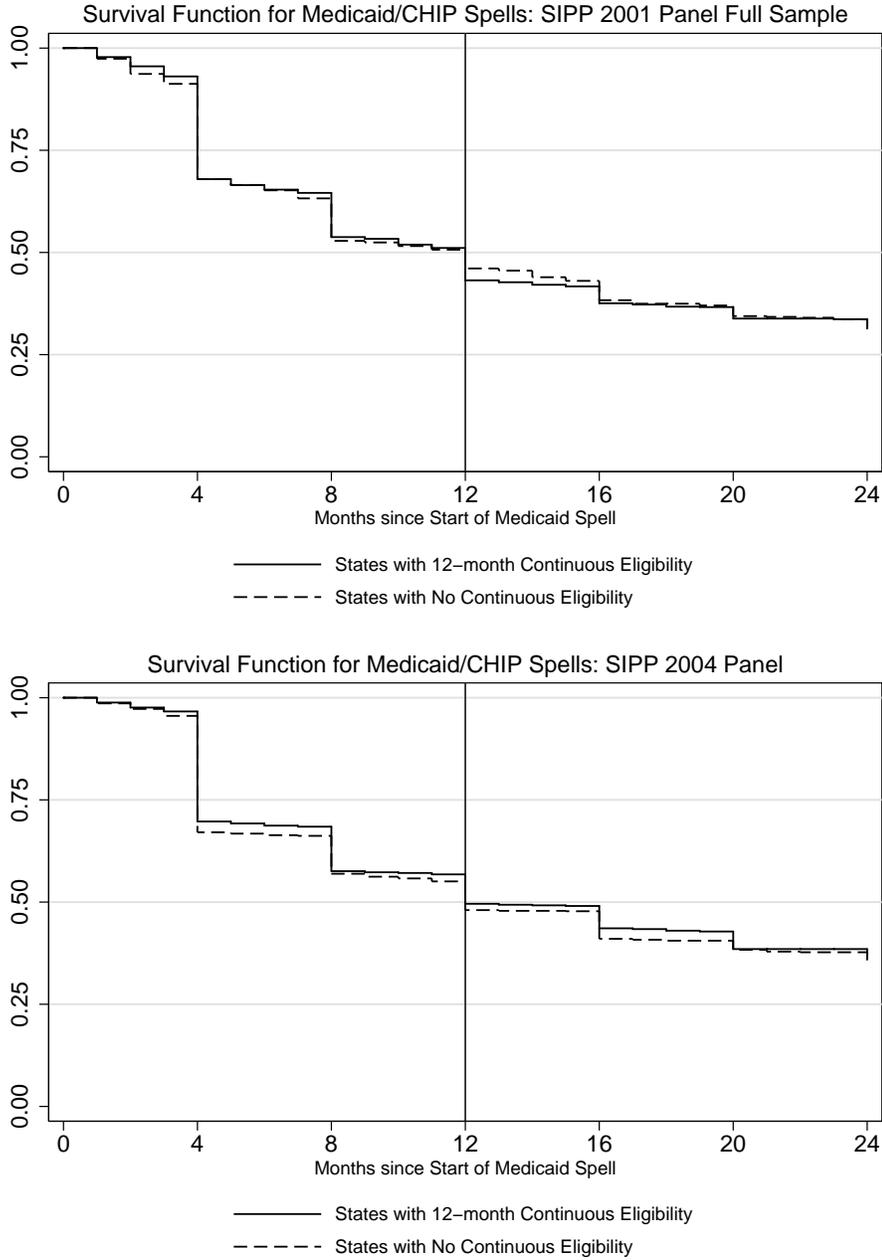|	Empirical Measures of Income		Empirical Measures of Income		Simulated Measures of Income			Simulated Measures of Income		
	Truncated Full Sample		Truncated Long Gap Sample		No Income Effect ($\rho=0$)			Income Effect ($\rho=1$)		
Month in Spell	Point Estimates	Upper 95% CI	Point Estimates	Upper 95% CI	e=0.5	e=0.2	e=0.05	e=0.5	e=0.2	e=0.05
-5	65	357	160	764	912	534	374	511	371	270
-4	28	302	69	697	912	534	374	511	371	270
-3	-48	93	-25	322	912	534	374	511	371	270
-2	-40	92	9	346	912	534	374	511	371	270
-1	-76	29	-21	124	912	534	374	511	371	270
0	0	0	0	0	0	0	0	0	0	0
1	33	230	-204	51	912	534	374	439	336	260
2	16	251	-294	-6	912	534	374	439	336	260
3	-8	206	-416	-76	912	534	374	439	336	260
4	-3	218	-488	-116	912	534	374	439	336	260
5	-183	20	-548	-63	912	534	374	439	336	260
6	-259	-54	-596	-90	912	534	374	439	336	260

Notes: Model calibration assuming continuous labor supply choice. Empirical estimates are based on specification (11) where the point estimates trace out average family income around the start of a child's public insurance spell. Model parameters are calibrated using 2004 SIPP panel income data, federal income tax rates, published Medicaid/CHIP eligibility cutoffs and CHIP spending data. Note that the implied Hicksian labor supply elasticity when $\rho = 1$ is $\frac{e}{\rho e+1}$, which is 0.33, 0.16 and 0.048 for e =0.5, 0.2 and 0.05 respectively.

Figure 1: Survival Functions for Medicaid/CHIP Spells

Survival Function for Medicaid/CHIP Spells: SIPP 2001 Panel Full Sample

Months since Start of Medicaid Spell

——— States with 12–month Continuous Eligibility
– – – – – States with No Continuous Eligibility

Survival Function for Medicaid/CHIP Spells: SIPP 2004 Panel

Months since Start of Medicaid Spell

——— States with 12–month Continuous Eligibility
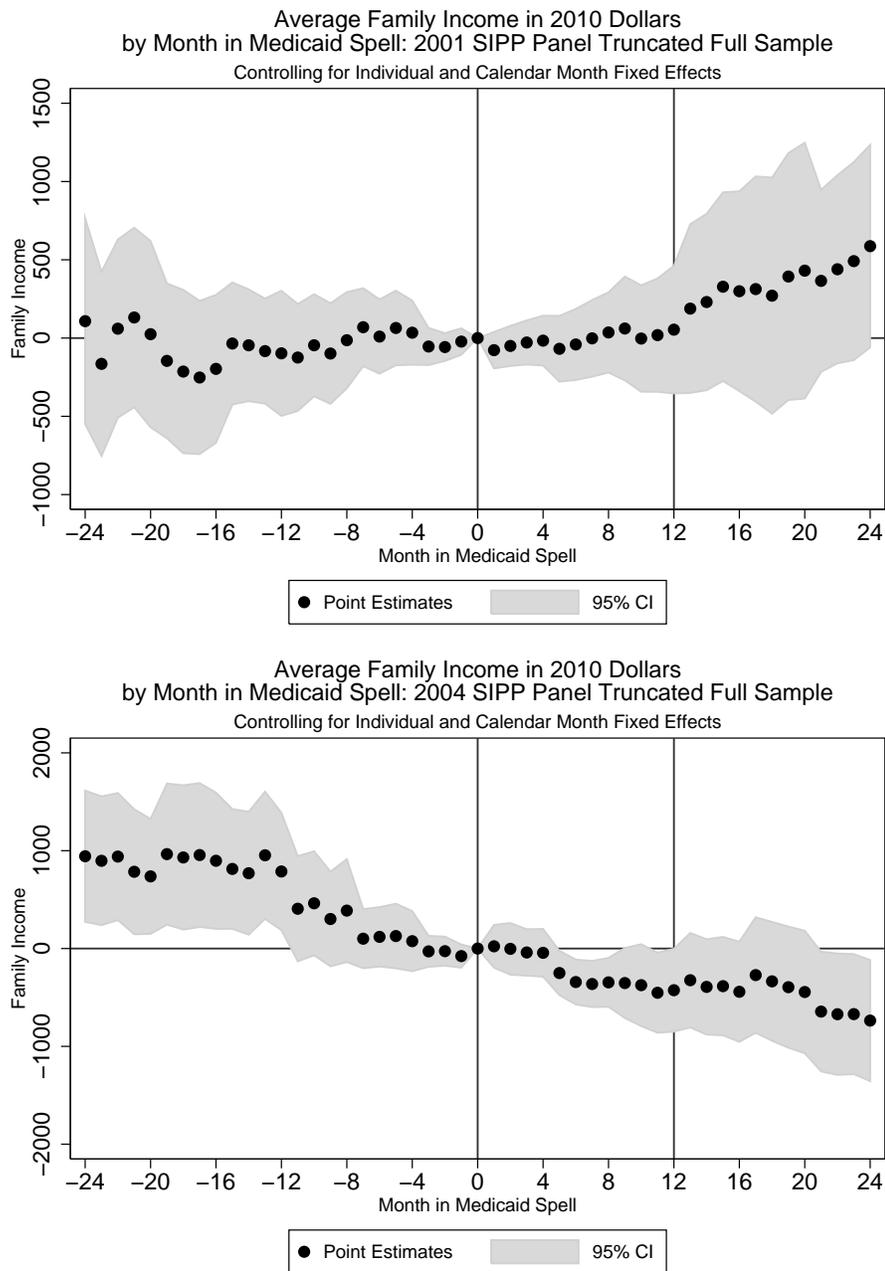– – – – – States with No Continuous Eligibility

Notes: Plotted are survival functions for children's public insurance spells for the SIPP 2001 and 2004 panels. The dashed and solid lines represent Medicaid/CHIP spells for children in states that provide 12 months of continuous eligibility and those that do not. The sample in the 2001 panel includes 3096 fresh spells for 2310 children from 1257 sample units in states with continuous eligibility and 2270 fresh spells for 1742 children from 934 sample units in states not providing continuous eligibility. The 2004 panel sample includes 3312 fresh spells for 2843 children from 1642 sample units in states with continuous eligibility and 2535 fresh spells for 2217 children from 1255 sample units in states not providing continuous eligibility.

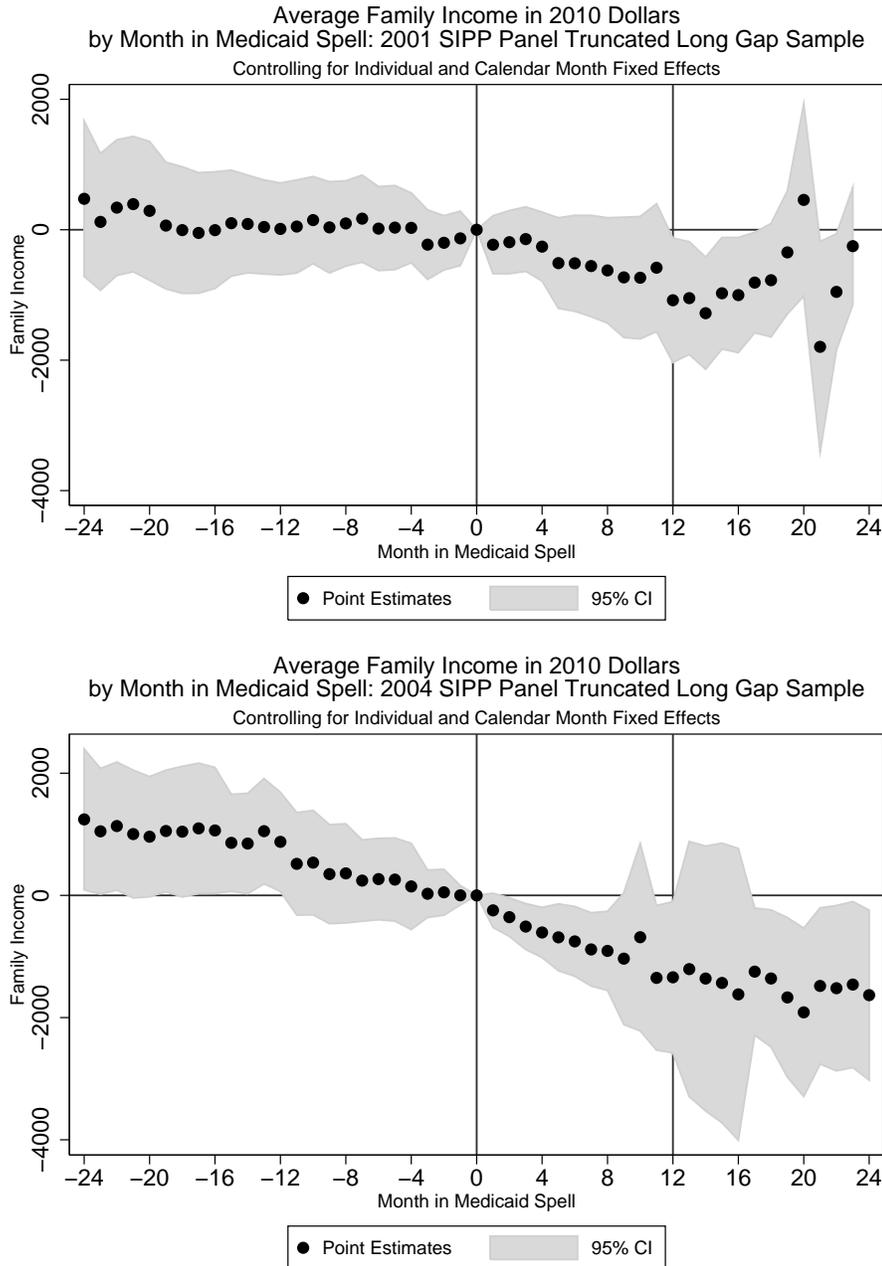Figure 2: Survival Functions for Medicaid/CHIP Spells: Long Gap Sample
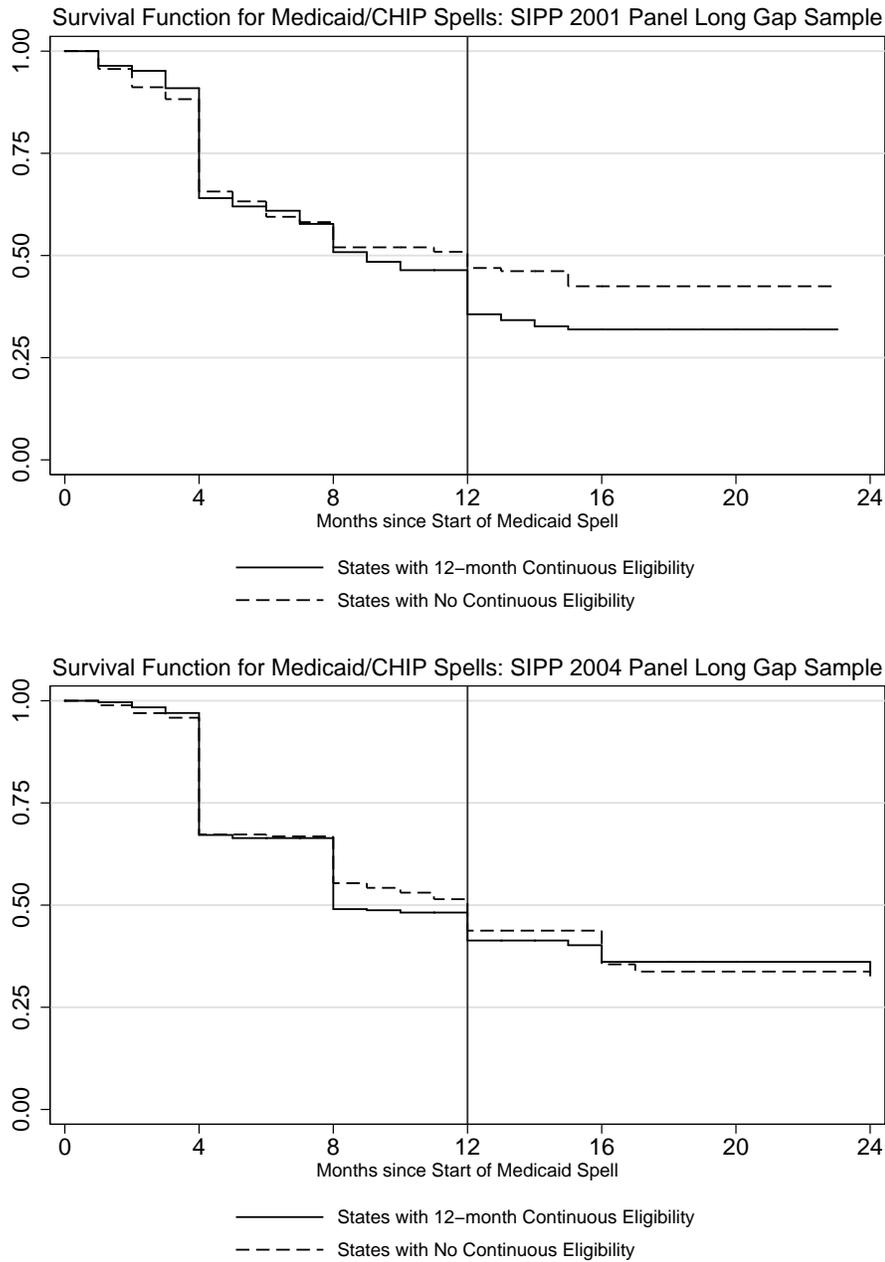
Survival Function for Medicaid/CHIP Spells: SIPP 2001 Panel Long Gap Sample

Months since Start of Medicaid Spell

——— States with 12–month Continuous Eligibility
– – – – States with No Continuous Eligibility

Survival Function for Medicaid/CHIP Spells: SIPP 2004 Panel Long Gap Sample

Months since Start of Medicaid Spell

——— States with 12–month Continuous Eligibility
– – – – States with No Continuous Eligibility

Notes: Plotted are survival functions for children's public insurance spells for the SIPP 2001 and 2004 panels. The dashed and solid lines represent Medicaid/CHIP spells for children in states that provide 12 months of continuous eligibility and those that do not. The sample includes children who were not covered by public insurance for 12 months before the start of a spell. The sample in the 2001 panel includes 501 fresh spells for 501 children from 302 sample units in states with continuous eligibility and 375 fresh spells for 373 children from 220 sample unites in states without continuous eligibility. The 2004 panel sample includes 815 fresh spells for 804 children from 494 sample units in states with continuous eligibility and 659 fresh spells for 656 children from 386 sample unites in states not providing continuous eligibility.

Figure 3: Average Family Income by Month in Medicaid Spell: Truncated Full Sample

**Average Family Income in 2010 Dollars
by Month in Medicaid Spell: 2001 SIPP Panel Truncated Full Sample**

Controlling for Individual and Calendar Month Fixed Effects

● Point Estimates   95% CI

**Average Family Income in 2010 Dollars
by Month in Medicaid Spell: 2004 SIPP Panel Truncated Full Sample**

Controlling for Individual and Calendar Month Fixed Effects

● Point Estimates   95% CI

Notes: Plotted are coefficients from regressions of family monthly income in 2010 dollars on a set of indicator variables for months before or since the start of a fresh Medicaid/CHIP spell. The top panel plots the coefficients from the SIPP 2001 panel and the bottom plots those from the 2004 panel. The regression includes individual and calendar month fixed effects. The standard error is clustered at the SIPP sampling unit level. Month 0 (the month right before the start of a fresh Medicaid/CHIP spell) is the omitted category, and the corresponding coefficient is 0 by construction. The sample in the 2001 panel includes 3096 fresh spells for 2310 children from 1257 sample units, and the 2004 panel sample includes 3312 fresh spells for 2843 children from 1642 sample units. The sample is truncated by discarding the observations with Medicaid coverage before the start of a spell and those without Medicaid coverage after.
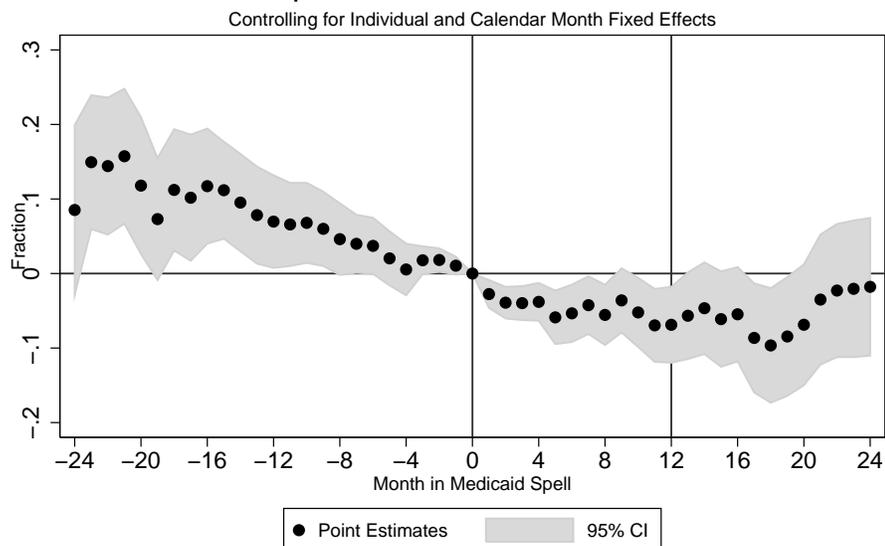
Figure 4: Average Family Income by Month in Medicaid Spell: Truncated Long Gap Sample

Average Family Income in 2010 Dollars
by Month in Medicaid Spell: 2001 SIPP Panel Truncated Long Gap Sample
Controlling for Individual and Calendar Month Fixed Effects

Month in Medicaid Spell

● Point Estimates   95% CI

Average Family Income in 2010 Dollars
by Month in Medicaid Spell: 2004 SIPP Panel Truncated Long Gap Sample
Controlling for Individual and Calendar Month Fixed Effects

Month in Medicaid Spell

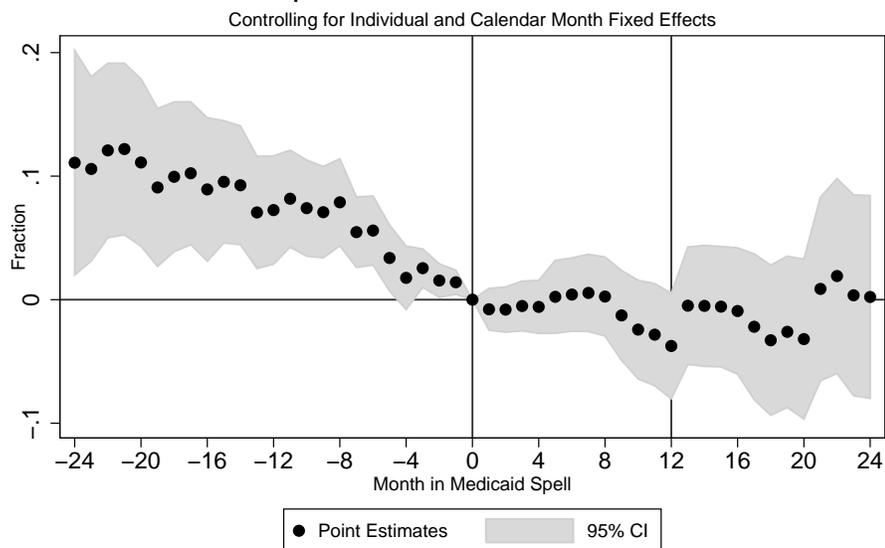● Point Estimates   95% CI

Notes: Plotted are coefficients from regressions of family monthly income in 2010 dollars on a set of indicator variables for months before or since the start of a fresh Medicaid/CHIP spell. The top panel plots the coefficients from the SIPP 2001 panel and the bottom plots those from the 2004 panel. The regression includes individual and calendar month fixed effects. The standard error is clustered at the SIPP sampling unit level. Month 0 (the month right before the start of a fresh Medicaid/CHIP spell) is the omitted category, and the corresponding coefficient is 0 by construction. The sample only includes children who were not covered by public insurance for 12 months before the start of a spell. The 2001 panel sample includes 501 fresh spells for 501 children from 302 sample units, and the 2004 panel sample includes 815 fresh spells for 804 children from 494 sample units. The sample is truncated by discarding the observations with Medicaid coverage before the start of a spell and those without Medicaid coverage after.

Figure 5: Fraction of Family Head Working Full Time in Medicaid Spell: Truncated Full Sample

Fraction of Family Heads Working Full Time All Weeks
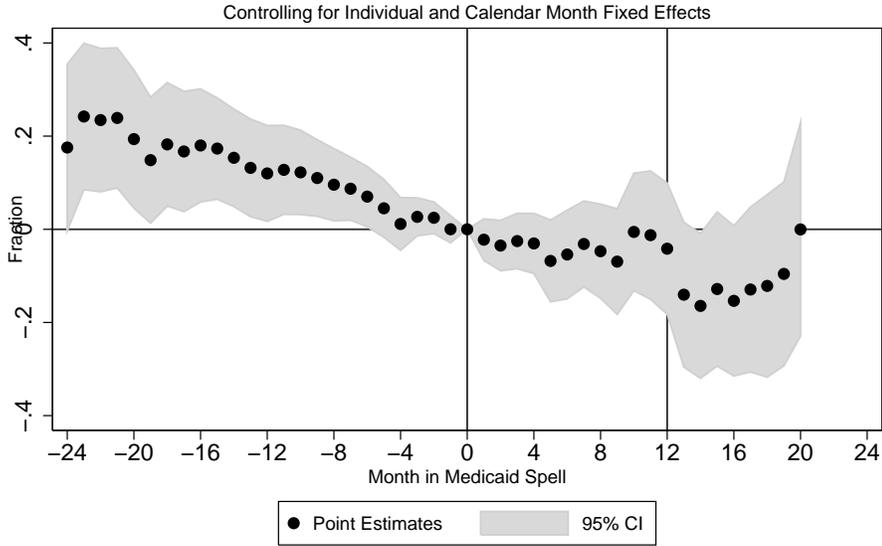### Month in Medicaid Spell: 2001 SIPP Panel Truncated Full Sample

Controlling for Individual and Calendar Month Fixed Effects

Fraction of Family Heads Working Full Time All Weeks
Month in Medicaid Spell: 2004 SIPP Panel Truncated Full Sample

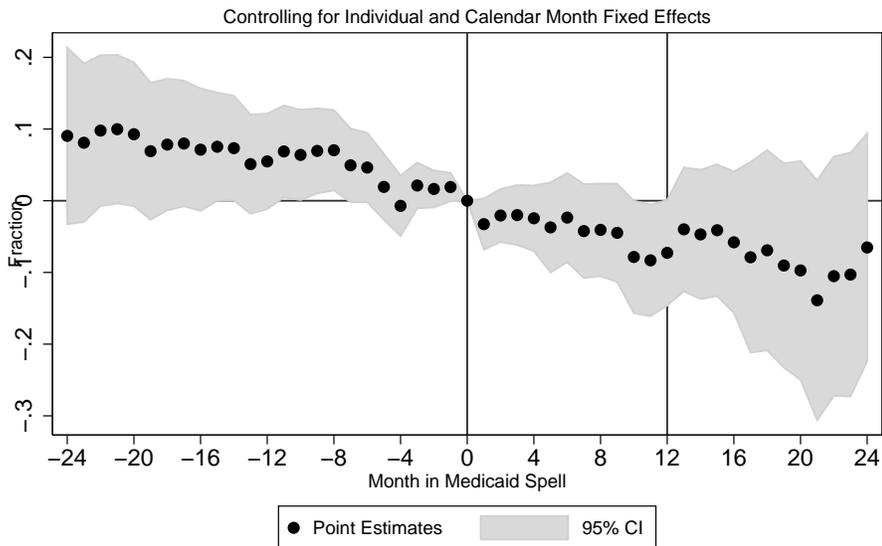Controlling for Individual and Calendar Month Fixed Effects

Notes: Plotted are coefficients from regressions of the indicator of whether the family head worked full time on a set of indicator variables for months before or since the start of a fresh Medicaid/CHIP spell. The top panel plots the coefficients from the SIPP 2001 panel and the bottom plots those from the 2004 panel. The regression includes individual and calendar month fixed effects. The standard error is clustered at the SIPP sampling unit level. Month 0 (the month right before the start of a fresh Medicaid/CHIP spell) is the omitted category, and the corresponding coefficient is 0 by construction. The sample in the 2001 panel includes 3096 fresh spells for 2310 children from 1257 sample units, and the 2004 panel sample includes 3312 fresh spells for 2843 children from 1642 sample units. The sample is truncated by discarding the observations with Medicaid coverage before the start of a spell and those without Medicaid coverage after.

Figure 6: Fraction of Family Head Working Full Time in Medicaid Spell: Truncated Long Gap Sample

Fraction of Family Heads Working Full Time All Weeks
Month in Medicaid Spell: 2001 SIPP Panel Truncated Long Gap Sample

Controlling for Individual and Calendar Month Fixed Effects

Fraction of Family Heads Working Full Time All Weeks
Month in Medicaid Spell: 2004 SIPP Panel Truncated Long Gap Sample

Controlling for Individual and Calendar Month Fixed Effects

Notes: Plotted are coefficients from regressions of the indicator of whether the family head worked full time on a set of indicator variables for months before or since the start of a fresh Medicaid/CHIP spell. The top panel plots the coefficients from the SIPP 2001 panel and the bottom plots those from the 2004 panel. The regression includes individual and calendar month fixed effects. The standard error is clustered at the SIPP sampling unit level. Month 0 (the month right before the start of a fresh Medicaid/CHIP spell) is the omitted category, and the corresponding coefficient is 0 by construction. The sample only includes children who were not covered by public insurance for 12 months before the start of a spell. The 2001 panel sample includes 501 fresh spells for 501 children from 302 sample units, and the 2004 panel sample includes 815 fresh spells for 804 children from 494 sample units. The sample is truncated by discarding the observations with Medicaid coverage before the start of a spell and those without Medicaid coverage after.