THE IMPACT OF CLAIMANT REPRESENTATION FEE SCHEDULES ON THE DISABILITY APPLICANT PROCESS AND RECIPIENT OUTCOMES

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Abstract

Many applicants to Social Security Disability Insurance (SSDI) retain representation to help with the approval process. The Social Security Administration imposes strict rules on representative compensation. Representatives are paid only if claimants are awarded disability, and are paid the lesser of 25 percent of claimant's past due benefits or a pre-specified maximum fee dollar amount ($6,000 since 2009). Because past due benefits are a function of the number of months claimants wait to be awarded, representatives face incentives to delay case resolution until past due benefits push representative fees past the maximum fee threshold. This paper uses difference-in-differences and bunching strategies to evaluate how these incentives impact SSDI applicant wait times. To do this, we use information on claimant characteristics and wait times constructed from the Disability Analysis File Public Use File (DAF PUF). Difference-in-differences estimation show that after changes in the maximum fee threshold, average wait times increase by 0.4-0.7 months among applicants for whom the fee threshold is more binding. We also observe bunching in the wait time distribution around the fee threshold kink in years after the policy change relative to the preceding years. Although both policy changes occur in years associated with economic recessions (2002 and 2009), we provide suggestive evidence that the increase in wait times is not driven by secular economic trends. Key limitations of this paper are data availability issues in the DAF PUF. In the DAF PUF, we are unable to observe who retains representations, so we can only observe the reduced form impacts of representative fee schedules on claimant outcomes. With more complete information on the claimant application and appeal process, we could more effectively probe the types of impacts that representation have on case outcomes and the sensitivity of the results to secular economic trends.

The paper found that:

- After the maximum representative fee threshold increases in 2002 and 2009, claimants for whom these fee thresholds were more likely to be binding initially observed increases in the number of months they waited between entitlement and award date on the order of 0.7 and 0.4 months respectively.
- After the maximum representative fee increases, estimated representative benefits also increase for these same claimants.

- There is also bunching or heaping in the number of months claimants wait from entitlement to award around the point where wait times would push 25 percent of past due benefits over the maximum representative fee threshold. This is the point where representative marginal revenue becomes zero.
- The policy changes occur in 2002 and 2009, which also happen to correspond to national economic recessions. We find similar patterns when we control for gender by education, by region-specific employment-to-population ratios, and claimant region and education. This would suggest the patterns are not driven by secular economic trends.

The policy implications of the findings are:

- The representative fee schedule impacts outcomes in the claimant application and appeal process. It may be beneficial to examine the incentives created by this schedule.
- In particular, increasing the maximum representative fee appears to increase wait times for claimants who were more likely to reach the old level of the threshold. This is consistent with representatives being able to increase revenue by allowing cases to draw out over longer periods of time.
- The analysis showing an increase in wait times for particular claimants when the fee cap is increased suggests that, by the same token, lower fee caps or flat fees could reduce wait times for claimants. However, this particular aspect of the policy should be interpreted in the context of the entire fee schedule and is, of course, subject to caveats that we discuss thoroughly in the paper. As disability representation is only paid when a claimant is awarded benefits, this is likely to impact the types of claimants whom representatives and attorneys choose to represent. Lower fee caps could change this selection, potentially resulting in equilibrium changes in who is represented and who obtains SSDI awards. Put differently, this means lower fee caps could lead attorneys to decline more cases and ultimately lead to fewer applications receiving awards.
- Because of data limitations, we are unable to examine decisions about who to represent or when to make representative fee agreements. With administrative SSDI applicant data, this is something we would be able to explore. We could also explore how representatives’ actions respond to the fee structure in more detail.

- From the bunching analysis, it is clear that representatives do not maintain perfect control over how long claimants wait for a case to be resolved. Rather than changing the fee schedule maximum fee cap, monitoring or regulating attorney action may be a way to ensure they do not draw out claimant wait times longer than necessary. Our research does not explore this avenue, and more work is needed to know how this could be effectively administered.

1 Introduction

The Social Security Disability Insurance (SSDI) program is one of the largest components of the US safety net in terms of dollars, transferring roughly \$11.7 billion a month to approximately 8.4 million disabled workers and their families (Social Security Administration, 2020). Since 2006, 2-2.8 million disabled workers apply for SSDI each year (Social Security Administration, 2020). The SSDI application process is complicated and drawn-out; in July 2019, 13.5 percent of 2018 claimants were still waiting a final decision and there were over still over 5,000 claimants from 2012 waiting for a final decision (Social Security Administration, 2020).

This process leads many claimants to obtain legal representation. In 2014, nearly 19 percent of all initial SSDI claims obtained representation, resulting in approximately \$1.1 billion of benefits being paid directly to representatives (Hoynes et al., 2016). The prevalence of representation has also become more common over time (see Figure 1). In 2000, the Social Security Administration (SSA) made approximately 15,000 representative payments per month. Today, that number has more than doubled to 30,000-40,000 payments per month. The SSA imposes a unique representation fee schedule, which restricts representative compensation to the lesser of 25 percent of the claimant's past due benefits or a prescribed amount, which has been \$6,000 since 2009. This schedule creates incentives for representatives to allow cases to continue unresolved, until 25 percent of expected past due benefits crosses the \$6,000 threshold.[1] After this point, the marginal revenue to the representative of further delays is zero. These incentives can potentially impact the amount of time it takes for applications and appeals to be resolved, and the amount of benefits an SSDI awardee will receive.

In this paper we use difference-in-differences and bunching strategies to determine if the representative fee structure leads to longer wait times or strategic bunching of cases

[1]We use the terms "fee threshold", "fee cap", and "fee ceiling" interchangeably throughout.

around the time when past due benefits would push representative fees past the maximum fee threshold. We use a novel dataset, the administrative Disability Analysis File (DAF) Public Use File (PUF), a 10 percent random sample of all SSDI recipients' benefit history between 1995 and 2015, which has been underutilized in existing work. Using difference-in-differences methods, we see if increases in the representative fee cap (in 2002 and 2009) lead to longer wait times for individuals with earnings-levels that make the fee cap relevant, relative to workers with earnings that make the cap less binding. We find that the increase in the cap in 2002 increases wait times for affected workers by about 0.7 months, on average. Similarly, the increase in the cap in 2009 increases wait times for affected workers by about 0.4 months, on average. Since only about 20 percent of claimants have representation, this would imply, under some assumptions, that wait times are 2-3.5 months longer for those with representation.

To explore the extent of precise, strategic behavior, we exploit changes in the maximum fee threshold over time and test for bunching around the kink. When comparing the distribution of wait times in the years before and after the 2002 and 2009 policy changes, there is significant heaping behavior leading up to the fee kink where 25 percent of past due benefits crosses the maximum fee threshold. Event study estimates suggest that this behavior begins around the time the policies are put in place. However, because the timing of both of these policy events correspond to economic recessions, it is possible these distributional shifts are due to aggregate economic trends which could influence claimant wait times. Both the difference-in-differences and bunching analysis is robust to controlling for gender by education by region specific employment to population ratios and including region and education controls.

Taken together, this evidence would suggest that the incentives created by the SSDI representative fee structure does lead representatives to slow down case processing, leading to higher fee payments and longer wait times. Because of the queuing and backlog for appealed cases, representatives do not seem to have precise control over how long it takes to process cases and for awards to be made.

Although there is significant work exploring the impacts of disability insurance programs like SSDI on claimant outcomes, such as labor supply (French and Song, 2014; Gelber et al., 2017; Maestas et al., 2013), financial stability (Deshpande et al., 2020), and household responses (Autor et al., 2019), there is very little work examining how the structure of SSDI affects other agents that participate in the process. We are only aware of one other paper that examines the role of SSDI representatives (Hoynes et al., 2016), which uses administrative SSA data to thoroughly document the descriptive patterns of who retains representation and how it correlates with case outcomes.

2 The SSDI Application Process and the Role of Representation

SSDI is a social insurance program, designed to insure against the risk that an individual becomes disabled and is no longer able to work and financially support themselves. To be eligible, workers must meet sufficient work requirements, meaning they have earned sufficient Social Security work credits in covered jobs and have sufficient work in the past 10 years. Workers must also document that they have a permanent disability that will last at least one year, keep them from performing the work they did previously, and prohibits them from adjusting to other work. Disabled workers who apply to SSDI will first have their application reviewed at a local SSA field office, to verify the individual meets all non-medical requirements. If they do, the application is passed on to a state-level Disability Determinations Services (DDS) agency. A DDS examiner will review the individual's case, evaluate available evidence, or seek additional evidence through a consultative examination. If it is determined that the worker is disabled, the case is returned to the field office where benefits are calculated and payments begin. Between 33 and 35 percent of initial applicants are allowed at this point (Social Security Administration, 2020). If the examiner determines the worker does not meet the disability requirements, the application is sent back to the field office, allowing the claimant to appeal the decision.

If a claimant chooses to appeal the decision, it is sent back to DDS to a new examiner for Reconsideration. Only 2-3 percent of initial applicants are allowed through Reconsideration (Social Security Administration, 2020). If the applicant is again denied, they can appeal the decision and it will then be sent to the appropriate hearing office, where it will be heard by an administrative law judge (ALJ). The ALJ will hold a hearing with the individual and any legal representation the individual has retained. The claimant can present further evidence regarding the severity of the disability, the onset date of the disability, or the limitations caused by the disability. Between 15 and 19 percent of initial applicants are allowed through an ALJ decision, meaning 31-36 percent of total recipients are approved by an ALJ (Social Security Administration, 2020).[2]

At any point in the process a claimant can retain legal representation. There is only limited data on the utilization of representation in SSDI appeals. Using SSA administrative records, Hoynes et al. (2016) find that about 13 percent of initial claims retained legal representation in 2010. By 2014, this had increased to 19 percent. Aggregated representation payout data suggests 377,054 representative payments were made in 2018, relative to 761,481 total awards (48 percent) (Social Security Administration, 2019). Claimants with disabilities related to back pain and major affective disorders are more likely to retain representation. On average over $1.1 billion a year of claimant benefits are paid directly to representatives for their services (Hoynes et al., 2016).

SSA has outlined two important conditions for SSDI representation compensation. SSDI representation (1) will only be compensated if the claimant is awarded disability insurance, and (2) will be paid a scheduled fee, which both the representative and the claimant must agree to, that does not exceed the lesser of 25 percent of past due benefits, or a prescribed dollar amount (Federal Register, 2009). This maximum amount has changed over time. The Omnibus Budget Reconciliation Act of 1990 specified the SSDI representative compensation formula, with a maximum prescribed amount of $4,000. This limit was increased to $5,300

[2]Because the hearing process take a long time, these are the estimates between 2005-2013, to avoid censored data.

on January 17, 2002, and then again to $6,000 on June 22, 2009 (announced January 29, 2009). Fee agreements that are approved by the SSA after these dates are eligible for the higher fee threshold.[3] Upon SSDI allowance, recipients are eligible for past due benefits from the month they became "entitled" for SSDI to the month they were awarded benefits. Because of the required five month waiting period, the entitlement date is typically five months after the established disability onset date. In partially favorable awards, where the judge establishes a later onset date, the entitlement date might not be five months after the disability onset date, but still corresponds to the first month the claimant is entitled to benefits. Throughout, we will refer to the first month the individual is entitled to benefits as the month of entitlement.

The conditional compensation stipulation incentivizes representatives to work with claimants that are likely to be awarded disability insurance and push for an award. If the goal of SSDI is to provide insurance payments to all eligible workers, then it is not clear that this incentive is misaligned. However, it could induce representatives to put more effort into cases that would otherwise be just below the margin of eligibility or cause them to select claimants who are more likely to have successful cases.

The fee structure stipulation may incentivize disability representatives to allow cases to move slowly through the decision process–which increases past due benefits and representative compensation–until 25 percent of past due benefits exceeds $6,000 (or the current maximum threshold). These incentives can result in claimants waiting longer for SSDI decisions and benefits, but could also increase the chance that claimants are awarded benefits and potentially the past due benefits they are owed. When considering welfare effects of the fee structure, longer wait times could result in claimants spending longer periods without income, potentially leading to even worse health. Although benefits during this waiting period would be eventually paid out, in the form of past due benefits, this might reduce welfare if claimants face binding borrowing constraints. Longer wait times that result in more past

[3]Since a claimant can seek representation at any point in the process, this date does not necessarily correspond to other dates in the SSDI process, such as the disability onset or application date.

due benefits and more dollars of representative compensation would also result in claimants paying a larger share of lifetime benefits to representatives. As past due benefits are also increasing in Average Indexed Monthly Earnings (AIME), incentives to delay case resolution would be stronger for low-income claimants, when it takes more months to reach the maximum fee limit. Anecdotally, there is some evidence of representative compensation rules influencing claimant outcomes. In a 2012 report (Social Security Advisory Board, 2012), field office and DDS employees allege that many professional representatives deliberately delay documentation and evidence to increase past due benefits and representative compensation.

3 Data

We use the Disability Analysis File (DAF) Public Use File (PUF) to examine how the SSA representative fee schedule affects claimant outcomes. The DAF is a compilation of SSA administrative data from various sources including the Mastery Beneficiary Record of beneficiary enrollment and the 831 files collected when the disability determination was made. It includes longitudinal records for adult SSI and SSDI recipients who have received benefits at any point since 1996, including spouses and dependents that receive benefits.[4] The PUF is a 10 percent random sample of beneficiaries in the DAF and has a more limited set of variables to avoid disclosure risk. For example, there is limited geographic information and dollar amounts are rounded.

There are two components to the DAF PUF. The first is a demographic file, which contains the cross-section of claimant attributes including things like date of birth, sex, primary and secondary disability diagnostic codes, and the claimants' application and entitlement date. The second component is a set of annual files, which include monthly records including claimant eligibility status, benefits due, and benefits paid. By combining payment and eligibility with the demographic file, we are able to construct the number of months

[4]Information for 10-17 year old recipients was added later, starting with beneficiaries that have received benefits after 2005.

between application and award and entitlement and award for each individual. Using the first recorded monthly benefits owed to the individual, we can also measure the individual's initial benefit level. These measures allows us to construct the claimants expected past due benefits by multiplying the initial benefit level by the number of months between entitlement and award.

There are limitations in the DAF PUF. Most relevant to the question at hand, there is no record of whether legal representation was retained. As such, we can only identify the reduced form effects of the fee schedule on claimant outcomes. However, we can see how effects differ for groups that are more or less likely to retain representation. In particular, we can examine effects across medical diagnostic groups. To maintain the anonymity and privacy of SSDI recipients, dollar amounts are also rounded. Dollar values between 1 and 7 were rounded to 4, values between 8 and 999 were rounded to the nearest 10, and values between 1,000 and 49,000 are rounded to the nearest 100. SSA reports that average monthly benefits for workers is $1,257.65 (in 2019), suggesting our measured benefit level could be up to 50 dollars off. This will lead to measurement error in our construction of past due benefits and representative fees. Dollar values are also top coded for the top half percent of positive values. As past due benefits are large relative to monthly payments, past due benefit payments that are made to claimants are likely to be muted by top coding. However, we can calculate the expected past due benefits because we can infer the entitlement month, award month, and initial monthly benefit. Dates are also reassigned to the 15th of the month for privacy. As such, we observe the number of months of past due benefits owed with error. However, as only full months are counted towards past due benefits, our measure should be within one month of the actual number of months benefits are owed. We also do not observe if an individual appeals a decision (and is thus more likely to retain representation).

We restrict the sample to only include individuals who applied in 1996 or later, so that we can accurately identify wait times. We also focus on workers and exclude spouses and dependents (who cannot be linked to workers).[5] We make several other restrictions to address anomalies in the data and mitigate the influence of outliers. We exclude people whose application to entitlement period is over 25 months, people whose application to allowance period is over 60 months, and whose initial benefit level was over $2,300. Combined, these restrictions exclude only three percent of claimants in the DAF PUF.

4 Empirical Strategy

We explore how the SSA representative fee schedule affects the appeal process and payments for claimants that are awarded disability. For each SSDI claimant, there is a kink in the potential representative fee, when the amount of time between the month of entitlement and allowance date pushes 25 percent of past due benefits over the current maximum fee threshold. The month of entitlement is typically five months after the established disability onset date, to account for the five-month waiting period. Prior to this kink, the representative's marginal revenue for one month of waiting is 25 percent of the individual's monthly benefit. Once the threshold is crossed, the marginal revenue for an additional month of wait time is zero. Because past due benefits are a function of both claimant AIME and how many months lapse between the month of entitlement and allowance dates (the month before effectuation), the amount of time it takes for 25 percent of past due benefits to exceed the threshold will vary for each individual.[6] We will first explore how changes in the maximum

[5]For fee agreement calculations, past-due benefits include the monthly benefits credited for all auxiliary beneficiaries unless they have appointed their own representative. We cannot connect dependents to workers and we do not know if dependents appointed their own representative. As such, our measures of past-due benefits may be under counted, but proportional, as both worker and auxiliary benefits are a function of the worker's earning history. In the combined DAF-PUF samples from 2015 and 2018, 92.3 percent of recipients are workers, meaning that this is a concern in less than 8 percent of cases (some workers are likely to have multiple dependents).

[6]Past due benefits and past due benefits payable to the claimant are different. For SSDI claimants, past due benefits are used to calculate representative fees, past due benefits payable are the benefits claimants receive after subtracting direct payments to representatives, reductions due to SSI receipt, and prior overpayments (`https://secure.ssa.gov/poms.nsf/lnx/0203920032`). Past due benefits are calculated differently for SSI claimants. Interim Assistance Reimbursement for state pre-payments is deducted before representative fees are calculated. Also, past due benefits include the month of effectuation. In this paper we focus on SSDI claimants and SSDI past due benefits.

fee threshold affect wait times for claimants that are more likely and less likely to face a binding fee threshold.

For an individual awarded disability insurance, monthly benefits are calculated using the same SSA Averaged Indexed Monthly Earnings (AIME) and Primary Insurance Amount (PIA) formulas used to calculate social security old-age and retirement benefits. The AIME is your average monthly earnings in your 35 years of highest earnings, after indexing for inflation. This number is then plugged into the progressive PIA formula with a 90 percent replacement rate for an AIME under $996 (2021 values), a 32 percent replacement rate for an AIME between $996 and $6,002, and a 15 percent replacement rate for an AIME over $6,002. Past due benefits is the number of full months between the entitlement date and award date, multiplied by the claimant's monthly benefit. As such, we can back out the number of months from entitlement to award it would take for 25 percent of past due benefits to meet the current maximum fee threshold by solving the following equation

$$Monthly\ Benefit_i * Months * 0.25 = Threshold \tag{1}$$

This yields the number of months to reach the threshold, $Months = \frac{4*Threshold}{Monthly\ Benefit_i}$. Because the monthly benefit varies across individuals this is an individual specific threshold, but the number of months to the threshold will fall as monthly benefits increase. For example, some individuals will reach the maximum fee in less than a year, while others with lower benefit amounts might not reach the threshold until after 3 years of waiting. Backlogs in the disability determination system result in different baseline propensities that the individual will wait long enough for the maximum representative fee will be reached. As seen in Figure 2, the threshold has increased twice, once in February 2002 by $1,300 (32.5 percent) and once in June 2009 by $700 (13.2 percent). An increase in the maximum fee threshold will lead to an increase in the number of months of past due benefits it takes to reach the fee threshold. In other words, representatives are compensated for more months of waiting.

Because representative fees are a function of past due benefits, and past due benefits are a function of monthly benefit levels, the maximum fee threshold will become more binding as monthly benefit levels increase. As a result, the probability of reaching the maximum fee threshold increases almost linearly with benefit level. That empirical relationship is consistent with differences in the probability of reaching the maximum fee being driven by differences in benefit levels, not necessarily differential selection on other dimensions that could affect wait time.

Although the relationship between benefit level and whether the fee ceiling binds is monotonic, the relationship between benefit level and whether the claimant is affected by the fee ceiling *increase* is not necessarily monotonic. As shown in Figure 2, claimants with back pay from $0 to $15,999 will not be affected by the cap increasing in 2002. Those claimants do not reach the cap when it is $4,000 or when it is $5,300. Similarly, although the cap increase in 2002 would increase the fee level for claimants with back pay greater than $21,199, the marginal fee for each additional dollar of back pay will remain zero, since these claimants will reach the cap under the $4,000 rule and under the $5,300 rule. The claimants who are affected are those for whom the cap binds prior to 2002 but, absent any changes in wait time, the new cap would not bind after 2002. Similar logic applies to the cap increase in 2009. In theory, this could induce a quadratic relationship between benefit level and whether the claimant is affected if claimants with high benefit levels have long enough wait times, on average.

We explore this empirically in Figure 3 by plotting the share of claimants in each $100 monthly benefit level bin who applied under the $4,000 cap and reached the $4,000 cap but who, based on their wait time, would not have reached the $5,300 cap. This provides a summary measure of the fraction of claimants at each benefit level who would be affected by the cap increase. Approximately 2.2 percent of claimants with benefit levels from $0-$500 would be affected by the increase in 2002 whereas about 7.1 percent of those with benefit levels above $500 would be affected. This figure exhibits a weak quadratic relationship, as fewer claimants with benefit levels above $1,700 would be affected by the increase than those with benefit levels from \$500-\$1,700. In Figure 4, we plot the share of claimants in each bin who applied under the \$5,300 cap and reached the \$5,300 cap but who, based on their wait time, would not have reached the \$6,000 cap. The figure exhibits a stronger quadratic relationship than Figure 3 but follows the same general pattern. Again, claimants with benefit levels from \$0-\$500 are unlikely to be affected by the cap increase (0.07 percent on average) and claimants with above \$500 are more likely to be affected (3.7 percent on average).

4.1 Strategy 1: Difference-in-Differences

Figures 3 and 4 motivate our difference-in-differences strategy. We combine this variation in the baseline propensity of reaching the maximum fee threshold driven by differences in earnings with policy changes in the maximum fee threshold. In particular, the relationship in both figures suggests that claimants with initial benefit levels between \$0-\$500 are less affected by the change in the fee ceiling than those with initial benefit levels above \$500. The intuition is that claimants with low benefit levels are unlikely to reach the fee ceiling in either period. As such, representatives have little incentive to change their behavior with regard to these cases after the threshold increases. Claimants with higher benefit levels, however, are likely to reach the fee ceiling prior to the increase but less likely to reach the fee ceiling once it's raised. As a result, representatives face a high return from lengthening those cases.

To formalize this, we categorize claimants with benefit levels from \$0-\$500 as the "unaffected" or "control" group and claimants with benefit levels above \$500 as the "affected" or "treated" group. We then estimate two difference-in-differences models comparing unaffected to affected claimants before versus after the fee ceiling increases in 2002 and 2009. Both models use the following specification:

$$Outcome_{it} = \alpha_0 + \beta(Affected_i * PostIncrease_t) + \gamma Affected_i + \phi PostIncrease_t + X_i + \varepsilon_{it} \quad (2)$$

When estimating the effect of the 2002 increase, we limit the sample to claimants who applied under the $4,000 cap or under the $5,300 cap. Then, we define $PostIncrease_t$ as equal to zero if the the claimant applied from January 1996 to February 2002 under the $4,000 cap and equal to one if they applied from March 2002 to June 2009 under the $5,300 cap. Similarly, when we estimate the effect of the 2009 increase, we limit the sample to claimants who applied under the $5,300 cap or under the $6,000 cap, and we define $PostIncrease_t$ accordingly. In both models, $Outcome_{it}$ is either the claimant's wait time (in months) from entitlement to approval or the hypothetical representative's fee based on the claimant's back pay and the fee ceiling in effect in year t. $Affected_i$ is equal to zero if the claimant's monthly benefit level is $0-$500 and equal to one if their benefit level is above $500. Finally, X_i is a matrix of individual-specific characteristics such as: sex, age, primary condition, and $100 benefit level bin fixed effects. Since Figures 3 and 4 show evidence of a quadratic relationship between benefit level and likelihood of being affected by the increase, we also explore robustness to excluding claimants with benefit levels above $1,500 or above $1,000.

We also estimate the following event study specification:

$$Outcome_{it} = \alpha_0 + \phi_y + \sum_{\tau=1997}^{2013} \beta_\tau Affected_i * \phi_\tau + X_i + \varepsilon_{it} \quad (3)$$

$Outcome_{it}$, $Affected_i$, and X_i are defined as they are above. ϕ_τ is a vector of year fixed effects from $\tau = 1997$ to 2013. β_τ then represents the average change in $Outcome_{it}$ for affected claimants associated with applying in year τ from 1997 to 2013, relative to the omitted year, 1996, and relative to the unaffected claimants in those years. If representatives are responding to the high return on lengthening cases that is specific to affected claimants, we should observe an increase in β_τ for years after 2002 and another increase for years after

2009, whether the outcome is wait time or representative fee. The key assumption with this approach is that affected and unaffected claimants would have followed similar trends if not for the fee ceiling increase. One testable implication of that assumption lies in the coefficients on β_τ from $\tau = 1997$ to 2009. If those coefficients are near zero from 1997 to 2001, that implies that affected and unaffected claimants were on similar trends since 1996 and only diverged once the fee was increased in early 2002. Similarly, if the coefficients from 2002 to 2009 exhibit a trend that is roughly flat, that again implies that affected and unaffected claimants were on similar trends prior to the second fee increase in mid-2009.

4.2 Strategy 2: Bunching

We will explore how claimant outcomes change in a close proximity around the kink point in the potential representative pay. This method will capture precise, strategic behavior that pushes wait times towards the fee threshold. Because some aspects of the application and appeal process are out of the representative's control, we might expect heaping rather than exact bunching. As see in Figure 2, marginal revenue for representatives will kink to zero at the point when the individual's past due benefits pushes representative fees over the threshold. Because the month of the kink varies with the monthly benefit amount, strategic bunching might not be detectable in the distribution of the raw number of months from entitlement to award (see Figure 7). This kink point will vary across individuals, by benefit level.[7] For this reason, we calculate the difference between the realized number of full months between entitlement and allowance date and the number of months that would reach the threshold. In other words we calculate

$$Month\ Waited\ Past\ Kink_i = Months\ Entitlement\ to\ Allowance_i - \frac{Fee\ Threshold_{yr} * 4}{Benefit\ Level_i} \quad (4)$$

[7]As seen in the second panel of Figure 8, the distribution of wait times is compressed towards zero for higher benefit claimants.

The fee threshold will either be \$4,000, \$5,300, or \$6,000 depending on the year, which is multiplied by 4 to capture the 25 percent rule. In Figure 8 we see that this distribution is centered around -12 months, but there is still considerable mass at 0, where the wait time between entitlement and allowance pushes 25 percent of past due benefits past the maximum fee threshold. There is no obvious evidence of strategic bunching.

To identify strategic bunching, we need to observe a counterfactual distribution. Rather than approximating a counterfactual distribution using smooth polynomials, we will exploit changes in the maximum fee threshold and use years prior to the threshold change as a counterfactual distribution. To determine how the kink in representative incentives affect claimant wait times, we will estimate how the density around the kink point associated with the 2002 and 2009 prescribed thresholds (\$5,300 and \$6,000 respectively) differs in 2002 and 2009 and in the years leading up to the policy change. In this case, we construct the Months Waited Past Kink measure in 2001 and 2002 (2008 and 2009) using \$5,300 (\$6,000) as the applicable Fee Threshold in equation (4). As seen in Figures 9 and 10, the year-to-year distributions are similar, but there is excess mass in the months leading up to and following the kink point in the latter year relative to the proceeding year.

To formally estimate these difference we collapse the data into one-month "Months Waited Past Kink" bins by application year. In each bin we sum the number of claimants in that bin in each application year. We then calculate the percent of claimants in each application year in each one-month bin, to facilitate a comparison across years, where the number of total claimants might fluctuate. We then estimate changes in the density within 60 months of reaching the kink in a regression format as follow

$$\begin{aligned} Percent\ of\ Claimants_{mt} = \sum_{\tau=-60}^{60} \beta_\tau(\tau\ months\ from\ kink_m) * PostIncrease_t \\ + \delta PostIncrease_t + \phi_m + X'_{mt}\Gamma + \varepsilon_{mt} \quad (5) \end{aligned}$$

The outcome of interest is the percent of claimants that applied in year t for whom the difference between realized wait time and the threshold was m months. Because the number of new claimants varies from year-to-year and we are comparing across years, we normalize by the total during the year rather than using the number of claimants. The main coefficients of interest are the β_τ, which trace out excess density in the post years, for each month-bin m relative to pre years. When examining the 2002 threshold increase we include application years 2000, 2001, and 2002. When examining the 2009 threshold increase we include application years 2007, 2008, and 2009. We limit the years to those right around the threshold, in hope that the clamaint cases will be the most similar, but include two pre-years to insure that estimated effects are not driven by changes in the counterfactual year.

We are interested in bunching right around the threshold where m equals zero, but include indicators for up to 60 months before and after the kink point to examine behavior throughout the distribution. All bins where the absolute value of m is less than 100 months are included (within 100 months of kink), so the β_τ coefficients are interpreted relative to the density changes in these regions of the distribution, over five years from the kink point. We include fixed effects for m, the number of month difference between realized wait time and the threshold, as well as average characteristics of claimants in the month/year bin, including sex, age, initial benefit level, and primary and secondary diagnostic code shares. The fixed effects capture the average density in the bin across all years in the panel, meaning the β_τ coefficients capture the relative change in density for claimants that filed in the post year. Robust standard errors are reported. The identifying assumption is that the density of wait times would have remained the same in the application year after the policy change if the threshold change had not occurred.

5 Results

5.1 Difference in Differences Results

Table 1 displays the estimated effects from equation (2). Column 1 excludes individual-specific controls, column 2 includes those controls, column 3 includes control but excludes claimants with benefit levels above $1,500 per month, and column 4 includes controls but excludes claimants with benefit levels above $1,000 per month. Turning to our preferred specification in column 2 of Panel A, we find that the fee ceiling increase in 2002 increased wait times by about three-quarters of a month, on average, for affected claimants. The estimated coefficient is smaller but not statistically distinguishable when we exclude controls in column 1. Columns 3-4 show that this result is robust to limiting the sample to claimants with more similar levels of income. Even comparing claimants with benefit levels from $0-$500 to those with benefit levels from $500-$1,000, we find stark differences in wait times after 2002.

Panel B of Table 1 displays the analogous results for the 2009 fee ceiling increase. Our preferred specification in column 2 of Panel B indicates that the fee ceiling increase in 2009 increased wait times by about half a month, on average, for affected claimants. As before, the estimated coefficient is qualitatively similar when excluding controls or when limiting the sample to claimants with more similar benefit levels. If we scale the coefficients in Panel A and Panel B by the amount of the fee ceiling increase, we recover estimates per dollar increased that are roughly similar in size. Table 2 follows the same structure as Table 1 except the outcome variable is the hypothetical representative's fee based on the claimant's benefit level, wait time, and the fee ceiling in effect at the time of the application. Consistent with increased wait times and higher likelihood of being affected by the fee ceiling increase, we see that estimated representative fees increase for claimants in our "affected" group relative to the unaffected group.

Figures 5 and 6 display the event study coefficients for these two outcomes. Figure 5 shows that affected and unaffected claimants followed similar trends in wait time from 1996-2001. In 2002, however, when the fee ceiling increases, we see an increase in wait time specifically for affected claimants. From 2002-2009, the trend in wait time is relatively flat, except for a divergence in 2006 and 2007. In 2009, when the fee ceiling increases again, we again see a larger increase in wait times for affected claimants than unaffected claimants. The estimated effects by year for representative fee follow a similar pattern.

These results suggest that SSDI representatives respond to changes in the fee structure that affect their payment. When the fee ceilings increase in 2002 and 2009, wait times increase specifically for claimants for whom the representative faces a high return to lengthening the case time. One concern with this strategy is that unaffected and affected claimants are fundamentally different groups of people since, by definition, unaffected claimants have lower incomes. The parallel pre-trends in Figures 5 and 6 partially address this concern since they highlight that despite their differences, these groups follow similar trends in wait time and in hypothetical representative fee prior to 2002. However, both fee ceiling increases occur shortly after recessions, and it remains possible that the unaffected and affected groups are differentially impacted by economic downturns. Tables 1 and 2 assuage this concern to some degree by showing that the estimates are robust to restricting the sample and comparing groups with more similar incomes.

To further test the possibility that these results are confounded by recessions, we re-estimate the relationship in equation (1), but control for the annual gender by education by region employment to population ratio (constructed using the March CPS) and for education and region fixed effects. Education is split into four groups (less than high school, high school, some college, and a college degree) while there are 10 regions that roughly correspond to census regions. As seen in Table 3, including labor market controls when looking at wait times lead to smaller point effects associated with the 2002 policy change and larger point estimates for the 2009 change, but these effects are still significant, implying an additional

0.3-0.6 months of wait time.[8] Impacts on estimates representative fees are not significantly different when we include labor market controls (see Table 4).

5.2 Distributional Bunching Results

We plot the estimated wait time density change in 2002 relative to 2000-2001 with 95-percent confidence intervals in Figure 11. We see a significant increase in density in the 12 months before and after the Months Wait Past Kink threshold. In some month bins, this represents a 0.2 percentage point increase in the number of claimants in a given monthly bin. In some cases, this represents a 20 percent increase in the density. The increase in density remains significant, although much smaller in magnitude, up through 24 months. There is also a significant reduction in density between 42 and 12 months prior to the kink threshold. We observe much smaller changes further away from the kink, most of which are not significant. A segment of claimants in 2002 waited substantially longer than similar claimants in 2000-2001, before the fee threshold increase. Congruent with the difference in differences results, this empirical pattern is consistent with representatives dragging out wait times to increase representative compensation.

When looking at the distribution around the 2009 threshold kink point there is also heaping, but it follows a slightly different pattern. As with the 2002 policy change, there is significantly more mass in the months just before the threshold, and significantly less mass in the proceeding months (42-18 months before the threshold). The magnitude of effects is about the same. However, unlike the 2002 policy change, we do not observe significantly more mass in the months following the kink point. The pattern turns precisely at the kink point and if anything there is less density past the kink. Although the empirical patterns differ slightly, they are both consistent with some claimants being pushed closer to the threshold

[8] If we separately estimate the effects for groups with above and below median employment to population ratios, the 2002 effects are larger for groups with higher employment to population ratios while the 2009 effects are similar across the two groups. If this pattern was driven by weaker labor market conditions impacting claimant selection and wait times we would expect larger effects in lower employment to population ratio cohorts.

and waiting additional months.

Because the bunching analysis reveals heaping rather than precise bunching, it is possible this is simply capturing a trending shift in the distribution over time. It would be concerning to see a similar shift in the distribution if we were to compare 2007 to 2005-2006 or any other combinations of years away from the policy change. We test this in an event study framework.

Using the collapsed month-by-year data, we estimate how the percent of claimants in the -12 to 12 month window of the 2002 kink threshold evolves over time since 1996 as follows

$$\begin{aligned} \textit{Percent of Claimants}_{mt} = \sum_{\tau=1997}^{2008} & \beta_\tau(-12 \leq m \leq 12) * (t = \tau) \\ & + \delta(-12 \leq m \leq 12) + \alpha_t + X'_{mt}\Gamma + \varepsilon_{mt} \quad (6) \end{aligned}$$

The outcome is the percent of total claimants in year t that fall in bin m where m is the number of months difference between entitlement to allowance wait time and the month the maximum fee threshold is reached. The main coefficients of interest are the β_τ which trace out across application years how the density changes for month bins within 12 months of the 2002 threshold relative to the other monthly bins. We control for the direct effect of being within 12 months as well as average characteristics of claimants in the cell and year fixed effects. The omitted reference year is 1996. Robust standard errors are reported.

We estimate a similar regression for the 2009 policy change, but look at changes in the density 0-18 months before the kink point as this is where the bunching is concentrated in 2009 relative to 2007-2008. This regression is as follows

$$\begin{aligned} \textit{Percent of Claimants}_{mt} = \sum_{\tau=2003}^{2013} & \beta_\tau(-18 \leq m \leq 0) * (t = \tau) \\ & + \delta(-18 \leq m \leq 0) + \alpha_t + X'_{mt}\Gamma + \varepsilon_{mt} \quad (7) \end{aligned}$$

In this specification data from application year 2002-2013 are included and 2002 is omitted as the reference year. In both cases we restrict the sample to exclude the other policy change. For many individuals, the policy changes lead to a difference that would cause overlap between the distance to the 2002 threshold and the 2009 threshold. This makes it impossible to interpret pre-trends prior to 2002 for the 2009 change and treatment effects after 2009 for the 2002 change.

These event study plots are provided in Figure 13. In both cases the pre-trends are flat, but begin to turn upward and become significant one to two years before the policy change. Since the policy change applies to all representative agreements approved after the policy date, and claimants can enter a representative agreement at any point in the application process, and agreements can include escalation clauses, it is possible, for example, that people who apply in 2000 or 2001 enter an agreement later and face the 2002 policy parameters. A slight pre-trend in the one to two year immediately proceeding the policy is not unexpected as we only observe application year, not agreement year (which is unknown).

For the 2002 policy change we see no changes in the density in the -12 to 12 month region of the distribution prior to 2000. In 2000 the point estimate is higher, but insignificant, but the effect is positive and significant in 2001. Then from 2002 on the effect is large and significant, consistent with approximately a 0.4 percentage point increase in the density per one month bin that gradually increases to about 0.9 percentage points.

For the 2009 policy change, we see no changes from 2002 to 2006, followed by an upward trend beginning in 2007, and continuing through 2009, the year of the change. This corresponds to a 0.5 percentage point increase, which increase to about 0.7 percentage points and then levels off.

Although the pre-trend is consistent with claimants entering representative agreements after their application year and becoming eligible, these periods also happen to loosely correspond to national economic recessions. There is evidence that during the Great Recession there was a large increase in disability insurance applications, appeals, and awards (Maestas et al., 2015). If these marginal claimants are selected differently than typical SSDI claimants, we might see an increased density of claimants at longer wait times (potentially around the kink threshold). We are already conditioning on average claimant characteristics in the one month bin, potentially accounting for some types of selection, but we cannot completely rule out that this is in part driven by aggregate secular trends. If we additionally control for the average employment to population ratio, and regional and education composition in each one month bin, the estimates are nearly identical (see Figure 14).

6 Implication for Disability Insurance and Future Extensions

These results provide suggestive evidence that SSDI representatives respond to fee structure incentives in ways that affect claimant wait time. The difference-in-differences analysis suggests that after the maximum representative fee threshold is increased, wait times increase for individuals at benefit levels where the fee threshold was more binding, relative to individuals at lower benefit levels where the fee threshold was less binding. The bunching results suggest that strategic behavior leads to heaping in the distribution of claimant wait times around the time the fee threshold increases.

Because we do not observe who retains representation, both sets of results should be interpreted as reduced form effects. Previous work estimates that in 2014, 19 percent of claimants retained representation (Hoynes et al., 2016). If we assume the share of claimants retaining representation did not respond to the maximum fee threshold (likely a strong assumption), the average treatment on the treated effect would be obtained by dividing effect sizes by the share of claimants that have representation. This would suggest that the 2002 fee increase was associated with a 3.6 month increase in wait times for claimants with representation while the 2009 increase was associated with a 2.1 month increase (assuming 19 percent take-up).

Both identification strategies rely on changes in the maximum fee threshold. Both fee increases occurred during economic recessions, which is potentially problematic if the recession differentially impacted the wait times of high- and low- benefit claimants or led to shifts in the wait time distribution. To the extent possible, we have shown that these estimates are robust to controlling for gender-by-education-specific labor market conditions in the region.

A richer data source could ameliorate these concerns. The DAF PUF does not provide any information on claimant representation. This information could be useful in several ways. First, examining patterns in the non-represented population could provide an informative falsification test. If only the wait times of claimants with representation adjust in 2002 and 2009 then there is less concern that these estimates are driven by secular economic trends. We could also compare the distribution of wait times for claimants with representation to that of claimants without representation in the same time period and place, to avoid secular patterns by holding economic conditions fixed. Second, more detailed information on when claimants seek representation and who represents them would allow us to explore alternative strategies. Using detailed information on representation agreement entry date, we could look at high frequency variation in close proximity to the maximum fee threshold changes. We could also examine whether the wait time effects are concentrated among certain representative types or groups. This data would also allow us to avoid secular concerns and make within period comparisons when looking at bunching.

With richer data, we could also explore how the fee structure's incentives impact other margins. Because the DAF PUF only includes claimants who are awarded disability, we are not able to examine how representative fees affect who retains representation or how this affects the probability of getting an award. Without clear information on the application and appeal process in the DAF PUF, we also cannot examine other behaviors, like representation delaying or withholding documentation to prolong the process.

Another relevant margin might be decisions about setting the disability onset date. The entitlement date depends on the disability onset date, as this establishes the earliest date a claimant is eligible to receive benefits. Establishing an earlier onset date would result in higher past due benefits, which will lead to higher representative payments as long as the maximum fee threshold has not been reached. It is possible representatives will put less effort into obtaining an earlier onset date when it does not increase their fees, even though it would still benefit the claimant. With data on representation and application and appeal details, we could explore how the likelihood of obtaining an earlier onset date changes when the wait time pushes representative fee payments over the maximum fee threshold. Much of this analysis would be feasible with the administrative 831 files or Mastery Beneficiary Record.

7 Conclusion

The SSDI application and appeal process is often drawn out and complicated. This leads many claimants to seek help from legal representation. The compensation rules of SSDI representation are dictated by the Social Security Administration to protect SSDI claimants. However, the structure of these rules may also create incentives for SSDI representatives to only take on cases that are likely to succeed and then to allow cases to drag on for long periods of time to increase past due benefits and representative compensation. In this paper, we provide suggestive evidence that the second feature of representative compensation rules leads to longer wait times for claimants.

In a difference-in-difference framework, claimants whose benefit-level make it more likely the maximum fee threshold is binding see longer wait times on average after the maximum fee cap is raised relative to claimants whose benefit-levels make the fee constraint less binding. This results in wait times that are about 0.5 months longer after the 2002 policy change and a net increase of about 1 month after the 2009 policy change.

There also appears to be strategic heaping in the wait time distribution around a fee kink threshold after the threshold changes relative to earlier years. These patterns tend to adjust in close proximity to the policy change, suggesting strategic behavior from legal representation leads to longer wait times. Although the timing of the policy changes correspond to national recessions, the estimates are robust to controlling for gender-by-education-specific employment conditions in the region. More work is needed to understand how this affects claimant outcomes.

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Tables and Figures

Table 1. Effect of Fee Ceiling Increases on Claimant Wait Time

	Wait Time from Entitlement to Approval (in Months)			
	(1)	(2)	(3)	(4)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009				
Affected x Post-2002	0.325***	0.735***	0.767***	1.131***
	(0.0652)	(0.0629)	(0.0631)	(0.0656)
Constant	13.20***	8.458***	8.188***	7.930***
	(0.0417)	(0.444)	(0.467)	(0.710)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.021	0.145	0.141	0.127
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	0.200**	0.440***	0.373***	0.236***
	(0.0806)	(0.0782)	(0.0791)	(0.0825)
Constant	15.84***	12.20***	11.77***	11.36***
	(0.0426)	(0.522)	(0.573)	(1.012)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.013	0.123	0.112	0.095
Controls	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** $p<0.01$, ** $p<0.05$, * $p<0.10$. Robust standard errors in parentheses. Panel A estimates the difference-in-differences model from equation (2) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. Panel B estimates the difference-in-differences model from equation (2) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. For both panels, column 1 excludes additional controls X_{it} and column 2 includes those controls (sex, age, primary condition, and benefit amount $100 bin fixed effects). Column 3 includes controls but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes controls but excludes claimants with benefit amounts above $1,000 per month. In this table, the outcome variable is the claimant's wait time from entitlement to claim approval (in months).

Table 2. Effect of Fee Ceiling Increases on Hypothetical Representative Fee

	Representative Fee (in $)			
	(1)	(2)	(3)	(4)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009				
Affected x Post-2002	598.4***	417.3***	419.1***	425.8***
	(6.502)	(6.259)	(6.275)	(6.584)
Constant	1,074***	1,009***	979.8***	1,045***
	(3.509)	(107.4)	(107.5)	(139.2)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.069	0.192	0.193	0.191
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	222.9***	137.3***	113.3***	60.98***
	(8.504)	(8.266)	(8.401)	(8.941)
Constant	1,290***	1,670***	1,558***	1,524***
	(3.885)	(122.1)	(126.2)	(199.8)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.032	0.143	0.143	0.144
Controls	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. Panel A estimates the difference-in-differences model from equation (2) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. Panel B estimates the difference-in-differences model from equation (2) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. For both panels, column 1 excludes additional controls X_{it} and column 2 includes those controls (sex, age, primary condition, and benefit amount $100 bin fixed effects). Column 3 includes controls but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes controls but excludes claimants with benefit amounts above $1,000 per month. In this table, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application.

Table 3. Effect of Fee Ceiling Increases on Claimant Wait Time, Labor Market Controls

	Wait Time from Entitlement to Approval (in Months)			
	(1)	(2)	(3)	(4)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009				
Affected x Post-2002	0.325***	0.326***	0.358***	0.771***
	(0.0652)	(0.0708)	(0.0710)	(0.0737)
Constant	13.20***	10.23***	9.069***	9.364***
	(0.0417)	(2.899)	(2.918)	(3.411)
Observations	1,124,297	946,908	868,335	651,131
R-squared	0.021	0.149	0.144	0.131
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	0.200**	0.556***	0.489***	0.350***
	(0.0806)	(0.0842)	(0.0851)	(0.0884)
Constant	15.84***	14.88***	13.37***	13.46***
	(0.0426)	(3.395)	(3.617)	(4.473)
Observations	1,145,400	1,023,424	884,025	605,915
R-squared	0.013	0.126	0.117	0.101
Controls	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. Panel A estimates the difference-in-differences model from equation (2) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. Panel B estimates the difference-in-differences model from equation (2) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. For both panels, column 1 excludes additional controls X_{it} and column 2 includes those controls (sex, age, primary condition, and benefit amount $100 bin fixed effects as well as the annual gender by education by region employment to population ratio, region, and education group fixed effects). Column 3 includes controls but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes controls but excludes claimants with benefit amounts above $1,000 per month. In this table, the outcome variable is the claimant's wait time from entitlement to claim approval (in months).

Table 4. Effect of Fee Ceiling Increases on Hypothetical Representative Fee, Labor Market Controls

	Representative Fee (in $)			
	(1)	(2)	(3)	(4)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009				
Affected x Post-2002	598.4***	396.3***	397.3***	407.7***
	(6.502)	(7.071)	(7.082)	(7.400)
Constant	1,074***	2,046***	1,652***	1,572**
	(3.509)	(591.6)	(560.2)	(622.9)
Observations	1,124,297	946,908	868,335	651,131
R-squared	0.069	0.190	0.190	0.189
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	222.9***	155.4***	127.4***	72.15***
	(8.504)	(8.892)	(9.023)	(9.561)
Constant	1,290***	3,104***	2,574***	2,413***
	(3.885)	(701.4)	(716.5)	(830.8)
Observations	1,145,400	1,023,424	884,025	605,915
R-squared	0.032	0.142	0.142	0.143
Controls	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. Panel A estimates the difference-in-differences model from equation (2) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. Panel B estimates the difference-in-differences model from equation (2) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. For both panels, column 1 excludes additional controls X_{it} and column 2 includes those controls (sex, age, primary condition, and benefit amount $100 bin fixed effects as well as the annual gender by education by region employment to population ratio, region, and education group fixed effects). Column 3 includes controls but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes controls but excludes claimants with benefit amounts above $1,000 per month. In this table, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application.

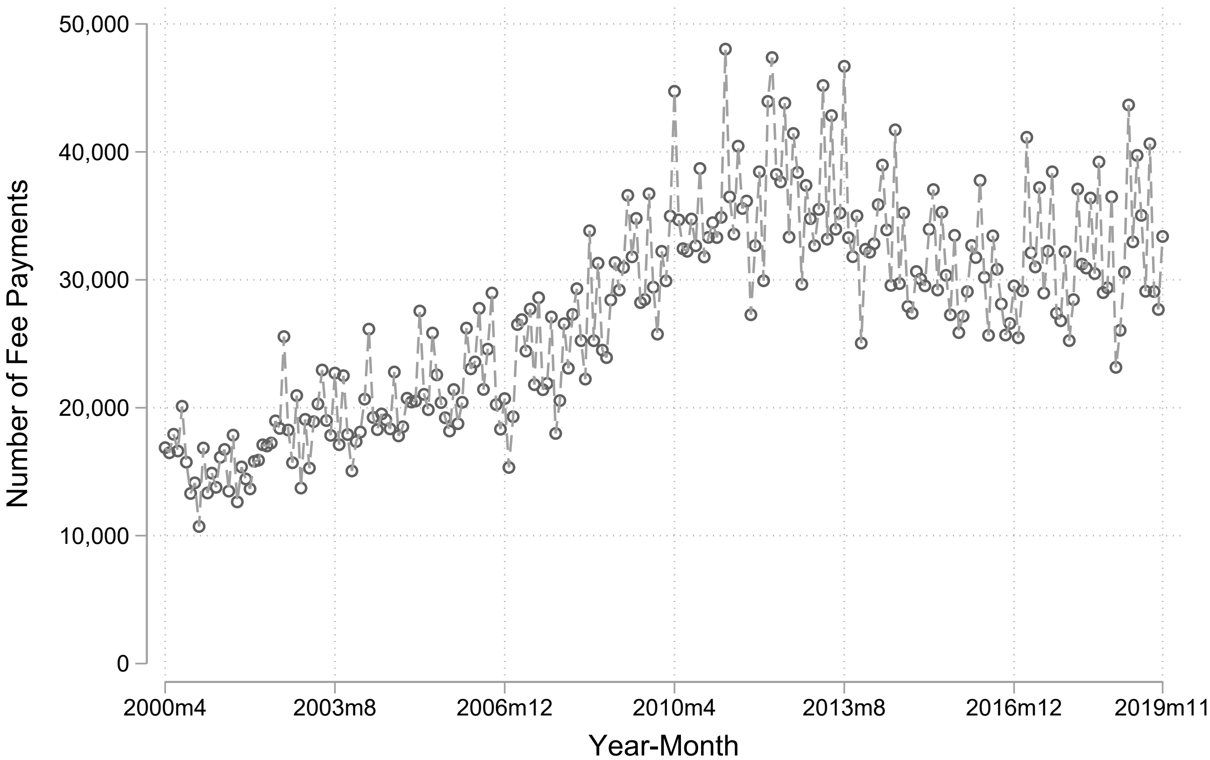

Figure 1: Number of Payments Made to Claimant Representatives by Month

Notes: The figure above plots the total number of fee payments made to claimant representatives by month. This amount has risen from 15,000 per month in 2000 to around 30,000 per month in 2019. The total amount paid has followed a similar pattern, growing from $40 million per month in 2000 to approximately $100 million per month in 2019. The total amount paid per fee agreement, on the other hand, rose substantially from 2000 to 2007, increasing from $2,250 to $3,500, but then fell from 2007 to 2014, bottoming out at around $2,800. As of 2019, the average amount paid per fee agreement is about $3,000.

Source: Authors' own calculations from SSA's "Statistics on Title II Direct Payments to Claimant Representatives"

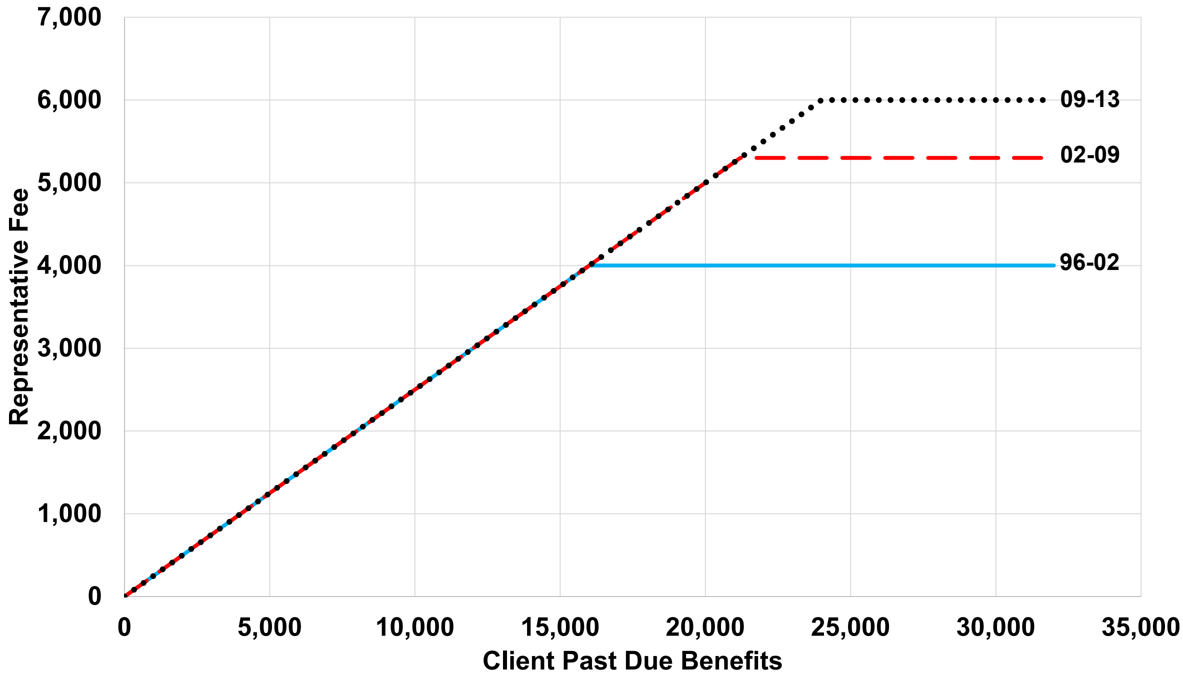

Figure 2: Representative Fee Schedule Over Time

Notes: Representative compensation schedule plotted over time. This structure was first adopted by the SSA in 1990, and a fee ceiling was set at \$4,000. The maximum fee was increased to \$5,300 on February 1, 2002 and to \$6,000 on June 22, 2009. Focusing on our years of interest: the schedule in effect from 1996-2002 is plotted with the solid blue line, the schedule in effect from 2002-2009 is plotted in the dashed red line, and the schedule in effect from 2009-2013 is plotted in the dotted black line. The maximum fee applies to agreements made between the claimant and representative after that date, regardless of their initial application date.

Source: Authors' own calculations.

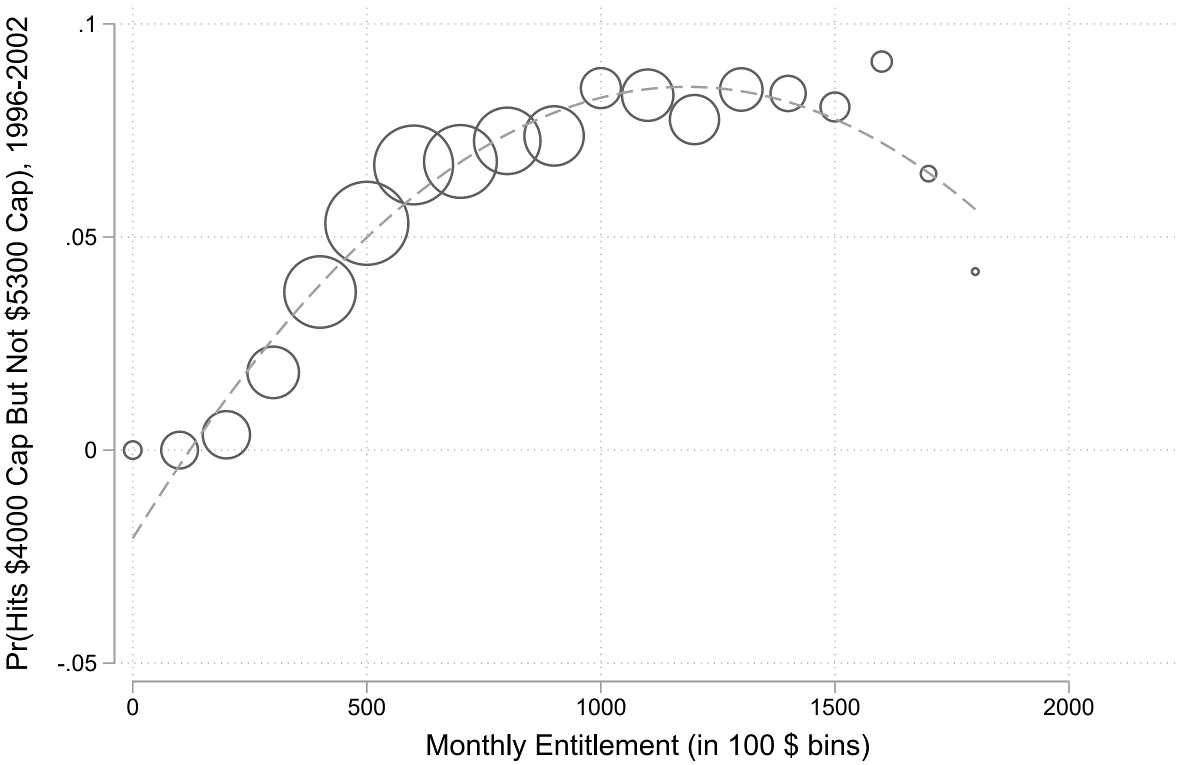

Figure 3: Share of Claimants from 1996-2002 Who Reached $4,000 Fee Ceiling but Would Not Have Reached $5,300 Fee Ceiling

Notes: Within 100 dollar monthly benefit bins we calculate the share of claimants for whom 25 percent of past due benefits exceeds the $4,000 fee ceiling but would not have exceeded the $5,300 fee ceiling. Past due benefits is the monthly benefit for the number of months the claimant is entitled benefits. This period is from the entitlement date to the award date. The fee regulation from the application year is used and only individuals who apply before February 1, 2002 are included. If the individual enters an agreement with representation at a later day, more generous compensation parameters might apply.

Source: Authors' own calculations from the SSA DAF-PUF.

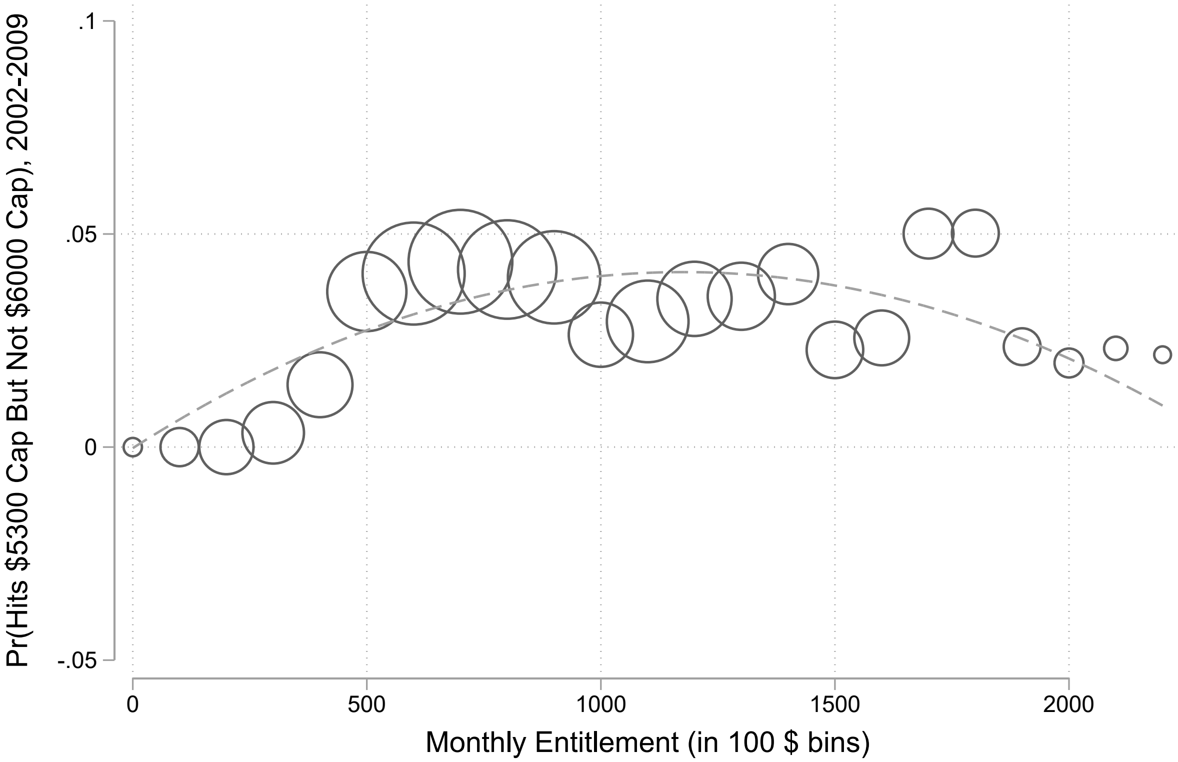

Figure 4: Share of Claimants from 2002-2009 Who Reached $5,300 Fee Ceiling but Would Not Have Reached $6,000 Fee Ceiling

Notes: Within 100 dollar monthly benefit bins we calculate the share of claimants for whom 25 percent of past due benefits exceeds the $5,300 fee ceiling but would not have exceeded the $6,000 fee ceiling. Past due benefits is the monthly benefit for the number of months the claimant is entitled benefits. This period is from the entitlement date to the award date. The fee regulation from the application year is used and only individuals who apply after February 1, 2002 but before June 22, 2009 are included. If the individual enters an agreement with representation at a different day, different compensation parameters might apply.

Source: Authors' own calculations from the SSA DAF-PUF.

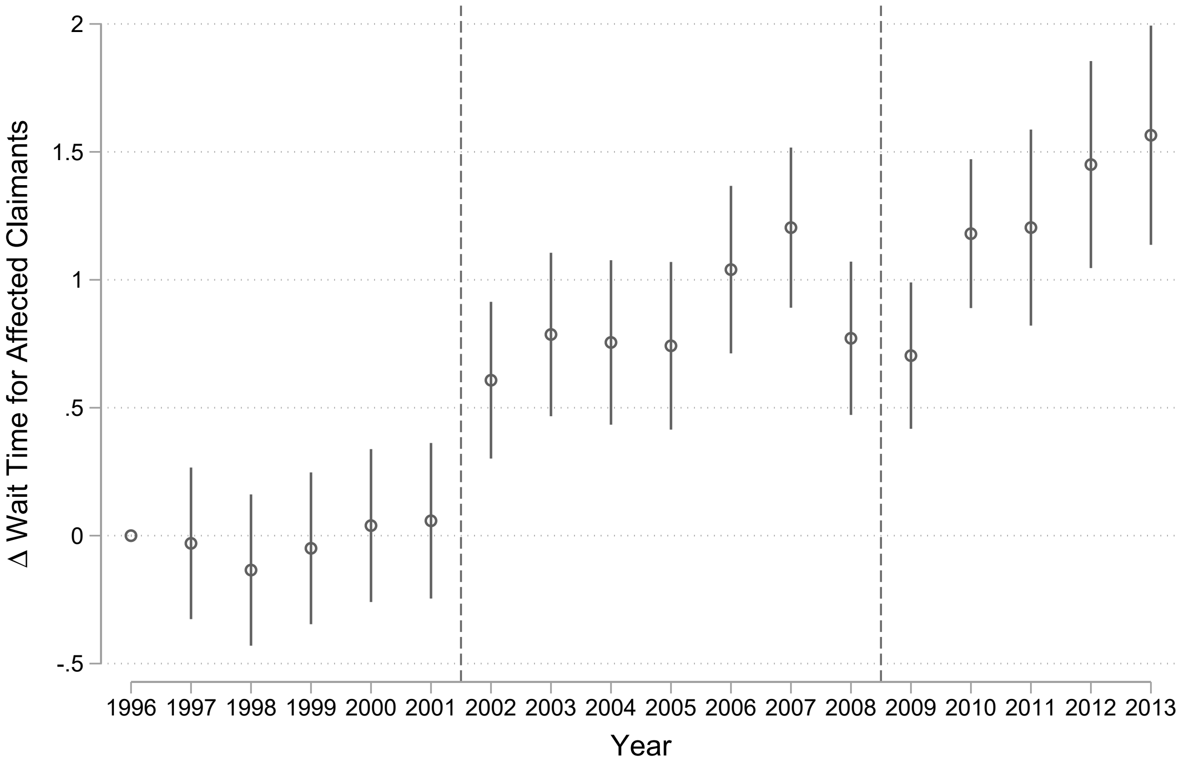

Figure 5: Change in Wait Time (in Months) For High Benefit Claimants
Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from eqn. (2). Each coefficient represents the change in wait time (entitlement date to approval date) in months since 1996 for claimants with above \$500 in monthly benefits relative to claimants with below \$500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, \$100 monthly benefit bin fixed effects, and year of application fixed effects are included. We include a dashed line after 2001 because the 2002 change occurs in February 2002, and we include a dashed line after 2008 because the 2009 change occurs in June 2009. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

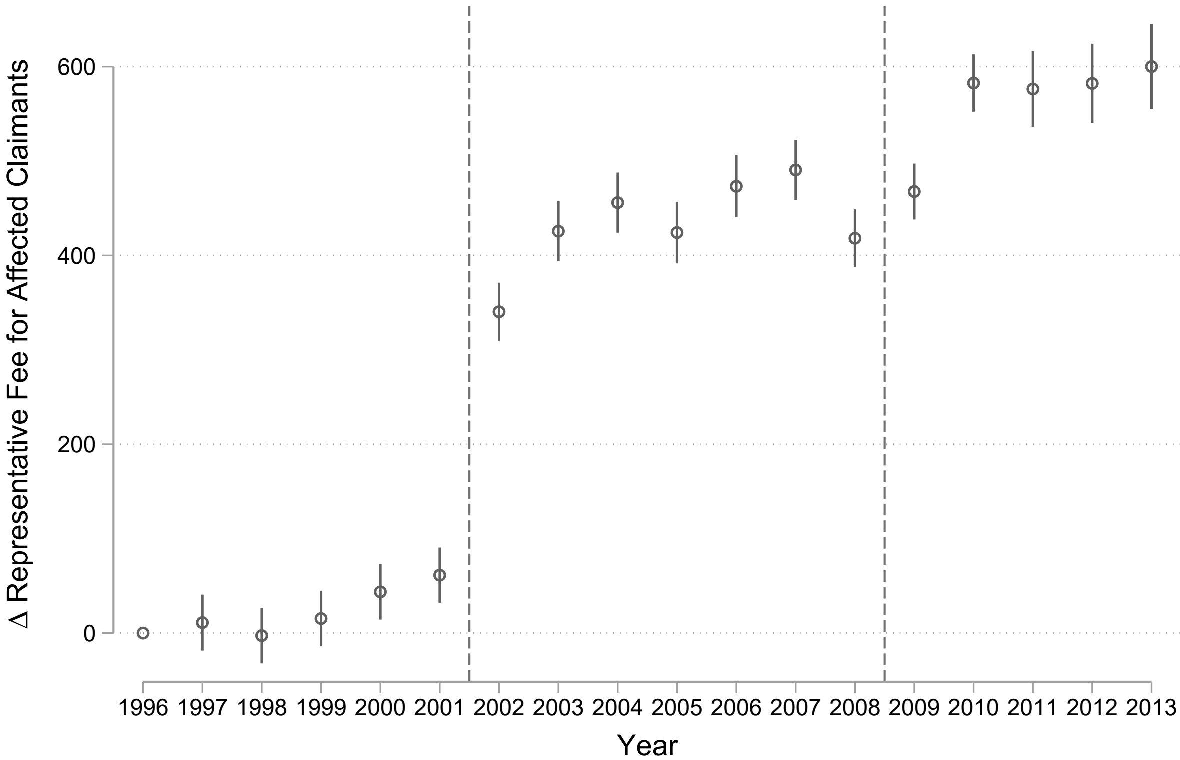

Figure 6: Change in Representative Fee (in Dollars) For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from eqn. (2). Each coefficient represents the change in representative fee in dollars since 1996 for claimants with above $500 in monthly benefits relative to claimants with below $500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, $100 monthly benefit bin fixed effects, and year of application fixed effects are included. We include a dashed line after 2001 because the 2002 change occurs in February 2002, and we include a dashed line after 2008 because the 2009 change occurs in June 2009. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

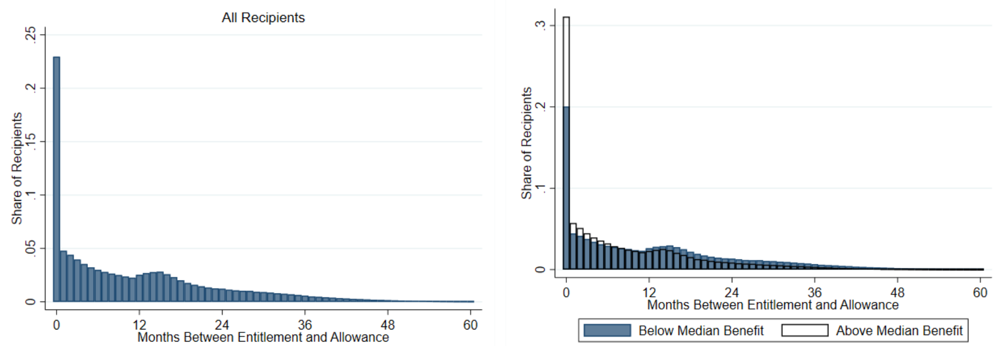

Figure 7: Distribution of Entitlement to Allowance "Wait Time" for Claimants

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date, applied between 1996-2013, and had 60 months or less of wait time between entitlement to award dates. The share of claimants whose entitlement to allowance time (in full months) is plotted in one month bins.

Source: Authors' own calculations from the SSA DAF-PUF.

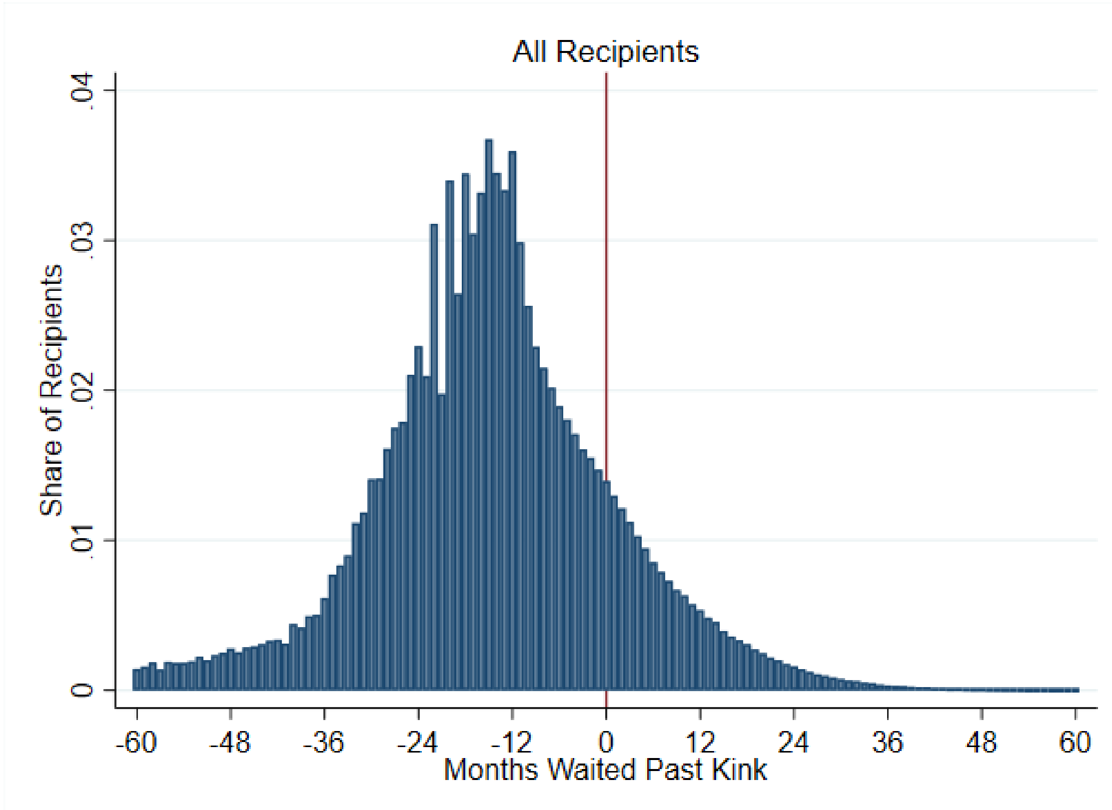

Figure 8: Distribution of Entitlement to Award “Wait Time” Relative to Maximum Representative Fee Threshold for Claimants that Likely Appealed

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date, applied between 1996-2013, and the differences between wait times and time to the representative fee kink was 60 months or less. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the year of application.

Source: Authors’ own calculations from the SSA DAF-PUF.

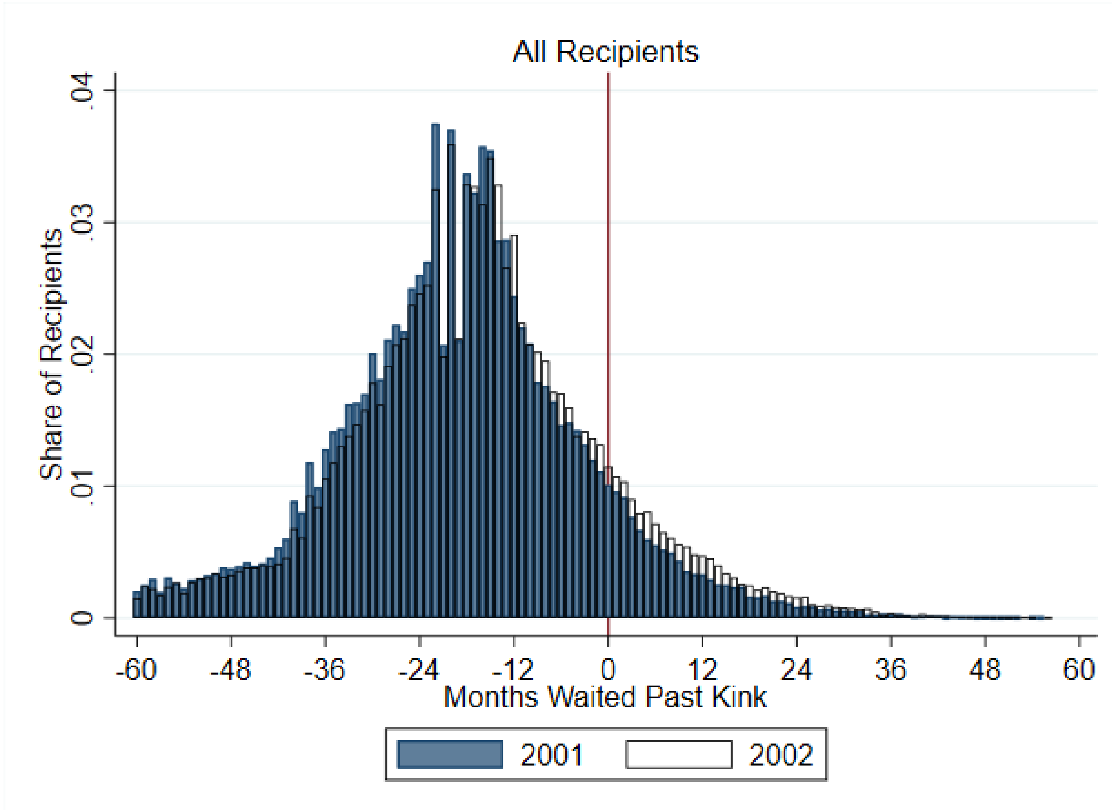

Figure 9: Distribution of Entitlement to Award “Wait Time” Relative to Maximum Representative Fee Threshold for Claimants in 2001 and 2002

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and the differences between wait times and time to the representative fee kink was 60 months or less. Claimants are plotted separately by those that applied in 2001 or 2002. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the application year.

Source: Authors’ own calculations from the SSA DAF-PUF.

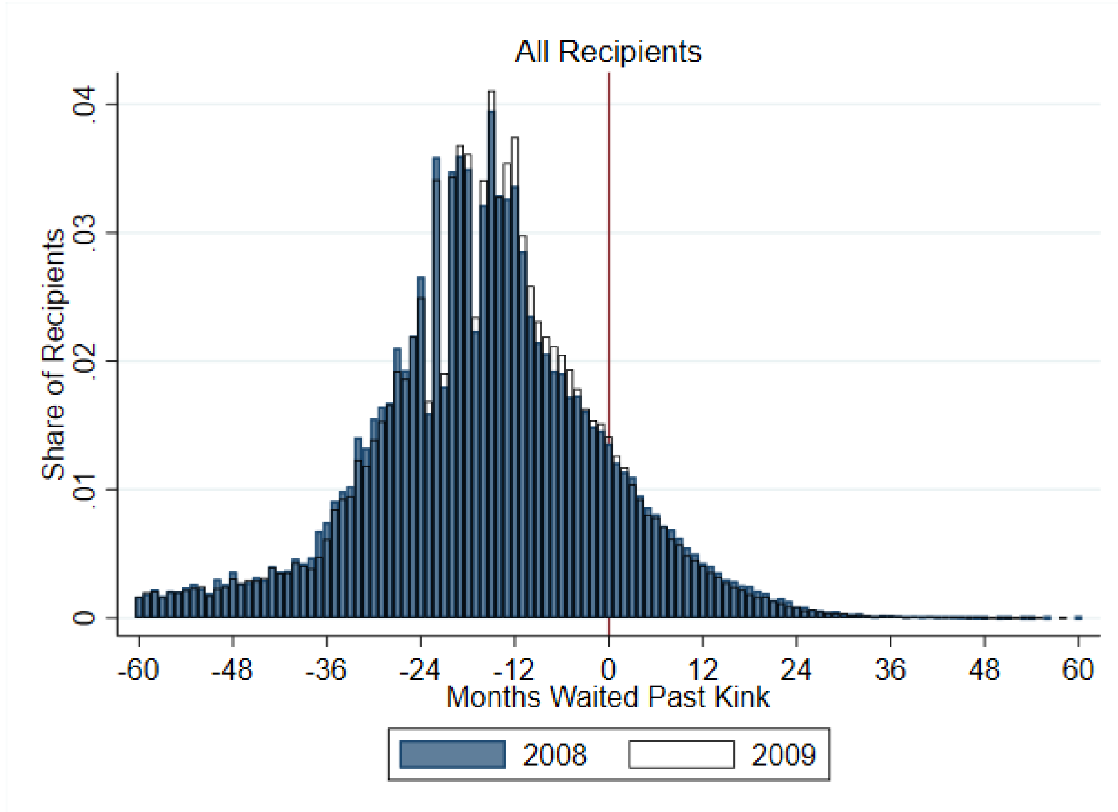

Figure 10: Distribution of Entitlement to Award “Wait Time” Relative to Maximum Representative Fee Threshold for Claimants in 2008 and 2009

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and the differences between wait times and time to the representative fee kink was 60 months or less. Claimants are plotted separately by those that applied in 2008 or 2009. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the application year.

Source: Author’s own calculations from the SSA DAF-PUF.

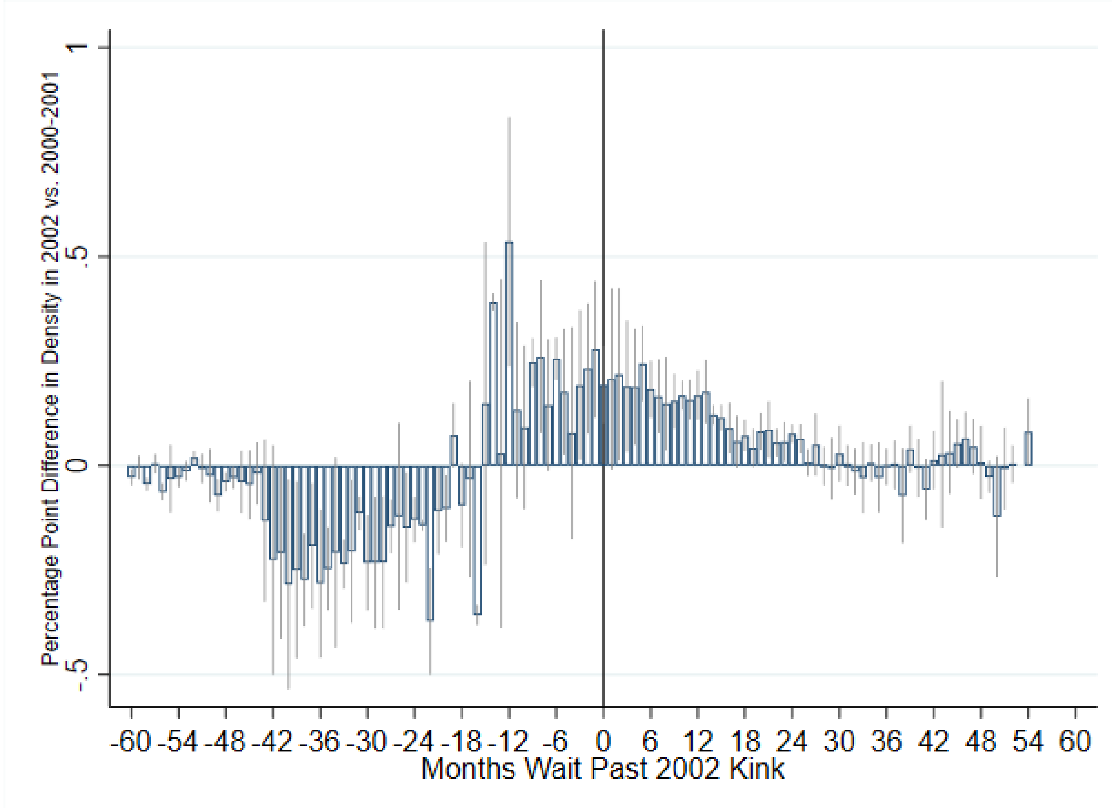

Figure 11: Difference in "Wait Time" Density Around 2002 Threshold in 2002 vs. 2000-2001

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date. Only claimants that applied in 2000, 2001, or 2002 are included. Each bar is the interaction coefficient from equation (4) with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial benefit level, and primary and secondary diagnostic code shares.

Source: Author's own calculations from the SSA DAF-PUF.

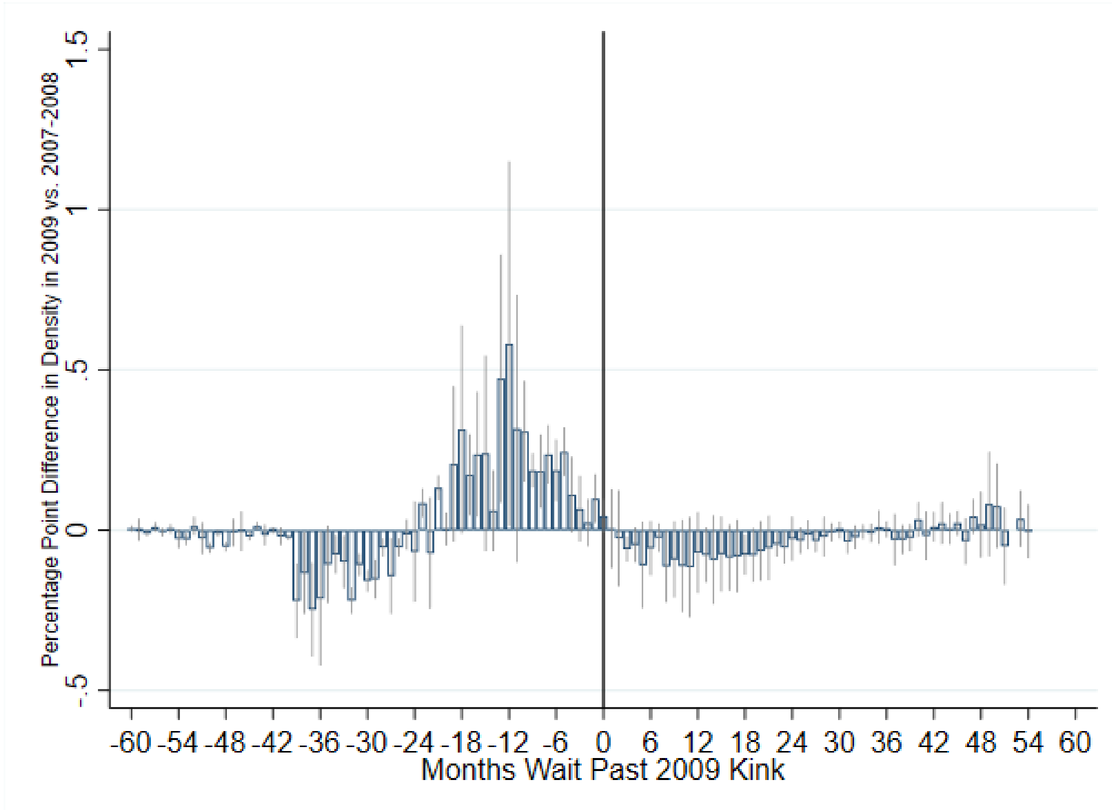

Figure 12: Difference in “Wait Time” Density Around 2009 Threshold in 2009 vs. 2007-2008

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date. Only claimants that applied in 2007, 2008, or 2009 are included. Each bar is the interaction coefficient from equation (4) with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial benefit level, and primary and secondary diagnostic code shares.

Source: Authors’ own calculations from the SSA DAF-PUF.

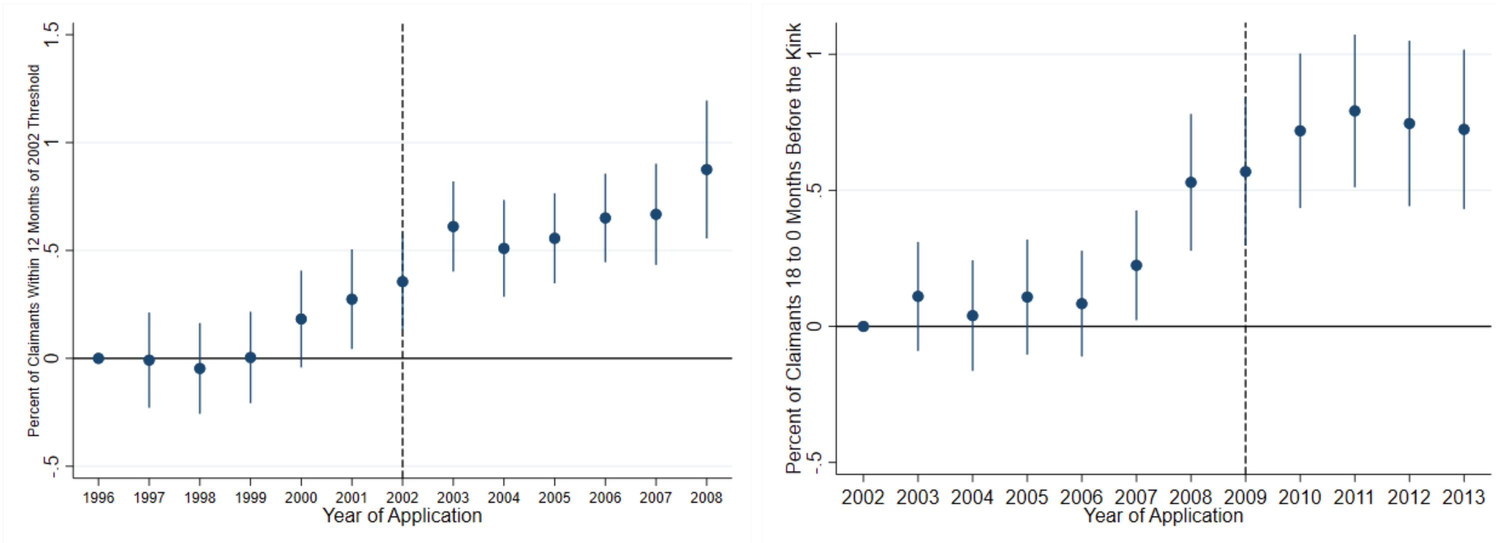

Figure 13: Event Study Difference in Bunching Around Maximum Fee Thresholds

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and were applied in 1996 or later. Individual level data is collapsed to month-from-the threshold bins. The coefficients represent the change in the share of claimants in the specified bins relative to other parts of the distribution in each year with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial benefit level, and primary and secondary diagnostic code shares.

Source: Authors' own calculations from the SSA DAF-PUF.

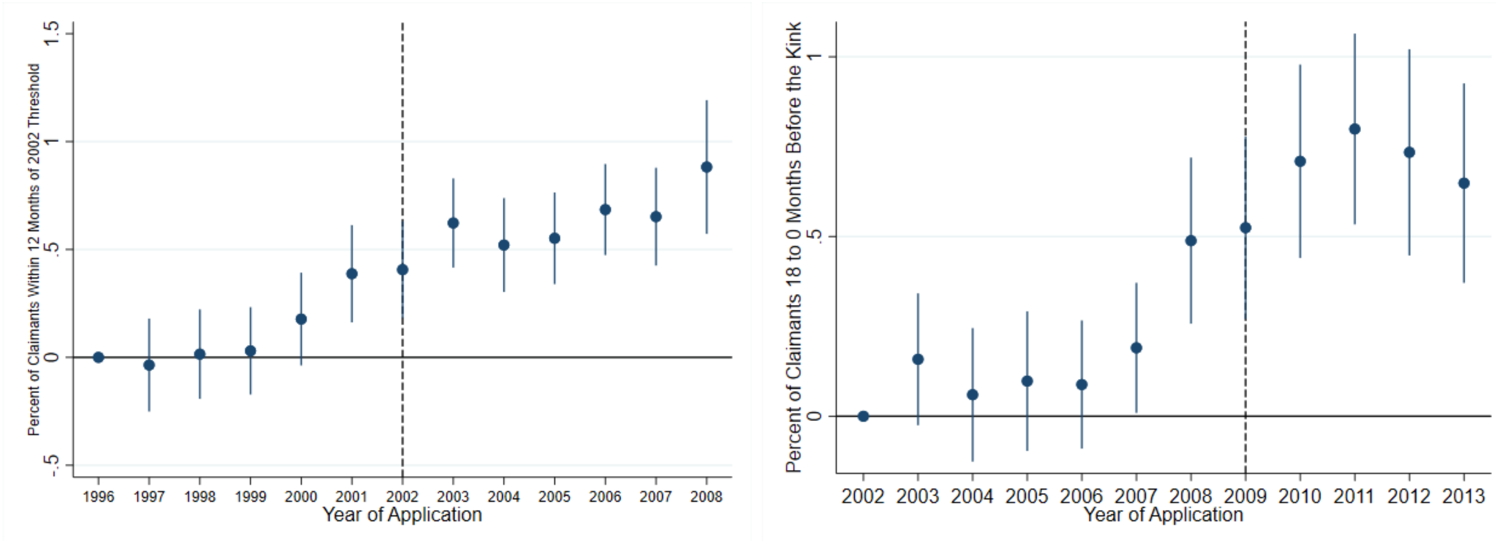

Figure 14: Event Study Difference in Bunching Around Maximum Fee Thresholds, Labor Market Controls

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and applied in 1996 or later. Individual level data is collapsed to month-from-the threshold bins. The coefficients represent the change in the share of claimants in the specified bins relative to other parts of the distribution in each year with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial benefit level, and primary and secondary diagnostic code shares as well as the average gender by education by region employment to population ratio in the bin and the share in each education and region bin.

Source: Authors' own calculations from the SSA DAF-PUF.

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