Representative Compensation and Disability Claimant Outcomes*

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Abstract

Many claimants of Social Security Disability Insurance (SSDI) retain legal representation to help with the approval process. The Social Security Administration imposes strict rules on representative compensation. Representatives are only paid if claimants are awarded disability, and they are paid the lesser of 25 percent of claimant's past due benefits or a pre-specified maximum fee ($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 the representative fees past the maximum fee threshold. We use difference-in-differences and bunching strategies to evaluate how these incentives impact SSDI claimant wait times. After changes in the maximum fee threshold, average wait times increase by 0.4-0.7 months among claimants 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 proceeding years. This indicates that the structure of representative compensation does matter for case outcomes, and highlights the importance of interactions with auxiliary agents so common in modern social programs.

Keywords: Social Security Disability, Attorney, Program Structure, Kink Points
JEL Codes: H5, J48, K23

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1 Introduction

A vast literature explores how the structure of various programs affects participants' incentives and behavior. However, program structure can also affect the incentives and behavior of administrators and intermediaries who are not direct beneficiaries of the program. There is far less work documenting how the incentives of these gatekeepers to social programs, such as caseworkers, judges, or legal representatives, affect beneficiaries' outcomes. These indirect effects can influence the overall effectiveness and efficiency of the program. As a case study, we document such patterns in one particular setting: legal representatives in the Social Security Disability Insurance (SSDI) program.

SSDI 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 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 15 percent of all initial SSDI claims obtained representation, resulting in approximately $1.1 billion of benefits being paid directly to representatives (Hoynes et al., 2021). 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.

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 fee ceiling, which has been $6,000 since 2009. Because past due benefits are increasing in the number of months claimants wait to be awarded, this schedule decreases representatives' incentives for timely case resolution. In other words, representative pay increases as wait

times increase 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 may impact the amount of time it takes for applications and appeals to be resolved, and by consequence, a host of claimant outcomes (Autor et al., 2017; Coe et al., 2014; Michaud et al., 2019).

In this paper we use two distinct identification strategies–a difference-in-differences design and a bunching analysis–to determine if the representative fee structure leads to longer wait times. To do this we use an underutilized dataset, the administrative Disability Analysis File (DAF) Public Use File (PUF), a 10 percent random sample of all SSDI recipients' benefit histories between 1995 and 2018. This allows us to explore detailed variation in claimant payments and wait times, in ways that are not possible with smaller surveys. Although the DAF PUF does not include information on representation, we exploit changes in the representative fee schedule to identify the reduced form effect on claimant wait times. 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 smaller cap increase in 2009 increases wait times for affected workers by about 0.4 months, on average. Based on existing estimates that about 15 percent of initial claimants have representation (Hoynes et al., 2021), this would imply, under some assumptions, that wait times increased by 2.5-4.5 months 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. However, because of the queuing and backlog for

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

appealed cases, representatives do not seem to have precise control over how long it takes to process cases and for awards to be made. As a result, we observe an increase in heaping around the fee kink rather than sharp bunching precisely above it. Because we do not observe who has a representative agreement, or the date the agreement was made, we can only determine timing relative to treatment with measurement error, but consistent with a causal channel, event study estimates suggest that this behavior begins around the time the policies are put in place.

One concern is that these distributional shifts in claimant wait time are due to the economic recessions in 2001 and 2009, not the 2002 and 2009 policy changes. We consider this possibility and provide suggestive evidence that the increase in wait times is not driven by such trends. First, both the difference-in-differences and bunching analyses are robust to controlling for gender by education by region specific employment to population ratios and to including region and education controls. Second, we estimate region-specific treatment effects and plot the relationship between those effects and region-level recession intensity. We do not find a relationship between how hard a region was hit by a recession and our estimated effect of the fee increase in that region. Moreover, we find lasting increases in wait times, which are more consistent with the permanent fee increase than transitory economic shocks.

To the extent possible, we also document potential mechanisms. Although the fee ceiling increase could affect both representative effort and attention to timeliness, we do not see systematic shifts in awardee characteristics or the propensity of early onset dates, which we might expect if representatives put in more effort, leading to higher award rates. As increased wait times are driven by actual delays, the pattern is more consistent with a reduced focus on timeliness rather than an increase in effort.

Taken together, this evidence suggests that the incentives created by the SSDI representative fee structure do lead representatives to slow down case processing, leading to higher fee payments for the representatives and longer wait times for the claimants. Wait times from

disability onset to benefit allowance are a critical piece of the disability claim process and are of recent interest to policymakers. As recently as 2019, several US senators introduced a bill specifically aimed at reducing SSDI wait times, citing the statistic that in 2017, more than 10,000 people died while waiting for their SSDI benefits to begin.[2] Autor et al. (2017) find that longer wait times lead to a reduction in post-decision labor supply, and Coe et al. (2014) show that while claimants wait for their SSDI decision they increase their reliance on SNAP benefits.[3] Other work on SSDI highlights that increasing the ordeal costs of applying decreases the targeting efficiency of the program, another potential consequence of long wait times (Deshpande and Li, 2019; Foote et al., 2018).

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., 2021). Hoynes et al. (2021) bring unique administrative SSA data to bear on the distinct and important question of whether having a representative improves case outcomes for claimants relative to not having a representative at all. Using a novel identification strategy that leverages geographic variation in disability representative markets, they find that retaining representation (versus not retaining representation) increases the probability of allowance in the initial review.

In addition to examining case outcomes, they also explore the impact of representation on various stages of processing time. They find that retaining representation is associated with a statistically insignificant 9 day increase in field office processing time (143% of the mean), a statistically insignificant 63 day increase in DDS processing time (68% of the mean) and a statistically significant 31 percentage point increase in the probability of onset to DDS

[2] S. 2496 (116th): Stop the Wait Act.

[3] Michaud et al. (2019) document a negative relationship between county-level wait times in one year and the number of applications in the next, suggesting that long wait times might discourage future applications.

decision taking longer than five months.[4] Despite these patterns, retaining representation significantly decreases total processing time on average by 316 days, largely by dissuading claimants with weaker cases from appealing SSA decisions. We explore a complementary question, how does a change in the representative's incentive structure affect claimant outcomes. Among successful claimants, representation might increase claimant wait times. Given the experience and expertise of disability representatives, it is possible retaining a representative improves some outcomes, but that the existing fee structure might lead to a smaller net benefit than alternative fee structures.

In addition to the specific insights we provide about legal representation in the SSDI system, we contribute to a separate literature that studies the connection between attorney compensation and client outcomes. Recent work on public defenders has shown that the structure of the compensation set by the public defender's office (flat fee vs hourly) significantly impacts pleas and convictions in criminal cases (Agan et al., 2021; Lee, 2021). Our paper adds new evidence by estimating the response of legal representation to compensation incentives in an important area of civil law that interacts with a crucial safety net program in the US.

Finally, this paper connects to an active policy debate about reducing the administrative hurdles and delays inherent to accessing social programs in the US. Researchers and policy analysts have examined these burdens in other domains, including Medicaid, unemployment benefits, and SNAP (Herd and Moynihan, 2020). Recent executive orders aim to curtail such hurdles for certain social security payments and disaster relief.[5] Ours is the first paper to document this phenomenon in the context of legal representation for disability application. We highlight how the structure of SSDI representative compensation can affect claimant wait times, a critical outcome on its own and a determinant of further downstream outcomes like future labor supply (Autor et al., 2017).

[4] As the authors note, this could be driven by representatives pushing for an earlier onset date, which increases past due benefits but does not affect the actual number of months the claimant waits.

[5] See: `https://www.bloomberg.com/news/articles/2021-12-13/biden-aims-to-slash-time-tax-for-social-security-storm-aid`

2 The SSDI Application Process

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 earn 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 (1) last at least one year, (2) keep them from performing the work they did previously, and (3) prohibit 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.

A claimant has 60 days to appeal a negative decision, in which case 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 again 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 all recipients

are approved by an ALJ (Social Security Administration, 2020).[6] If a claimant is denied by an ALJ, they can appeal again to the SSA Appeals Council then for Federal Review. Very few cases are pursued to these final stages.

3 Disability Representation Compensation and Incentives

At any point in the process a claimant can retain legal representation. Both attorneys and non-attorneys[7] can act as representatives. Estimates suggest that in recent years, around 83 percent of all cases with a representative have an attorney representative, but non-attorneys were not eligible to receive compensation until 2010, after the fee cap increases (Hoynes et al., 2021). There is only limited data on the utilization of representation in SSDI appeals. Using SSA administrative records, Hoynes et al. (2021) find that about 8 percent of initial claims retained legal representation in 2010. By 2014, this had increased to 15 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., 2021).

SSA has outlined two important conditions for SSDI representative compensation. SSDI representatives (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

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

[7]To be eligible, non-attorneys must hold a bachelor's degree, pass a written exam administered by SSA, obtain liability insurance, undergo a criminal background check, and complete continuing education courses (see Social Security Act, Section 206 [42 U.S.C.406].

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.[8]

Crucial to this compensation schedule are past due benefits. 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.[9] Throughout, we will refer to the first month the individual is entitled to benefits as the month of entitlement. Past due benefits are the claimant's monthly benefit level times the number of full months between the entitlement date and award date. 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. Thus, for pre-determined earnings history and benefit level, there is a number of months of "wait time" from entitlement to award date that will push 25% of past due benefits over the maximum fee threshold. At this point, the representative's fee is maxed out and no longer dependent on wait time or past due benefits.

To understand how this compensation structure affects the representative's incentives, it is useful to consider the representative's objective function. We assume the representative is profit maximizing. As such, the conditional compensation stipulation incentivizes representatives to take on claimants that are likely to be awarded disability insurance and to push

[8]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.

[9]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.

for an award. Given data constraints we are not able to examine this decision and will focus on the second feature of the payment structure.

The fee structure stipulation creates incentives for 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). Conditional on taking the case, the representative's expected revenue depends on both the probability of winning the case and the expected payout as follows

$$E(revenue_{it}) = \rho \min\{0.25 * b(X_i) * E(\,Wait\ Time_i), FeeCap_t\} \tag{1}$$

where ρ is the probability of winning; $b(X_i)$ is individual i's benefit level (which depends on their work history and other characteristics). Representative actions that increase wait time will increase expected revenue as long as wait time is below the prevailing maximum fee cap. Importantly, claimants and disability representatives do not have perfect control of the application timeline. For example, during the DDS review or ALJ appeal much of the wait time is driven by SSA action, not claimant or representative behavior. However, after both the initial review and Reconsideration, claimants are given 60 days (2 months) to appeal negative SSA decisions before the case is automatically closed. Claimants can use this time to gather more supporting evidence and decide if they want to appeal. This suggests there are several ways representatives could increase wait times.

First, they can put more effort into the case by collecting more evidence and building a more convincing argument, which might take time. Representatives might be willing to increase effort in this way if it increases ρ, the probability of winning the case. Second, they can put less focus on timeliness. For example, rather than working quickly to resubmit missing information or submit an appeal, they can de-prioritize the case, submitting it closer to the 60 day deadline. There are multiple steps when this could happen. This change in prioritization and timeliness need not be nefarious, but since there is not pressure to move the

case along quickly and in fact incentives to delay, representatives might move their focus to other demands. This reduction in timeliness is unlikely to affect the probability of winning, except in extreme circumstances where a deadline is missed and the case is closed. It can however increase wait times, past due benefits, and the representative's expected pay off. The reduction in timeliness is also likely to reduce psychic costs for the representative, as they relax their time constraints.[10]

Although both channels are potentially important, the monetary incentives associated with reducing timeliness will play a larger role in responses to the fee cap increase. If effort increases the probability of winning, the incentive to increase effort always exists. Proximity to the fee cap only marginally affects this since the payout for winning is large and positive, before and after the increase in the fee cap. However, the return to reducing timeliness jumps dramatically for cases near the pre-existing cap once the cap increases. We argue that representative responses to the change are more a result of this large increase in the returns to reducing timeliness rather than the small increase in returns to increasing the probability of winning.

For example, consider a simplified case where the individual's benefit level and expected wait time pushes past due benefits right to the pre-2002 cap of \$4,000. Suppose that the baseline probability of winning the case is 0.6, meaning the representatives expected revenue is $0.6 * 4,000 = 2,400$. Now, suppose the representative can exert additional effort and increase the probability of winning by 10 percentage points. This additional effort might include things like compiling a more complete documentation of the individual's disability and might take additional time. As such, suppose the additional effort also increases expected past due benefits to \$5,300. When exerting additional effort, the expected revenue is $0.7 * 4,000 = 2,800$ or \$400 dollars higher, since the \$4,000 cap is still in effect.

Now, consider what happens when the maximum fee cap is lifted to the 2002 level of \$5,300. Under this scenario, the representatives expected revenue when exerting additional

[10]This would suggest that reducing a focus on timeliness is both weakly revenue-increasing and cost-reducing.

effort increases to $0.7 * 5,300 = 3,710$, or \$1,310 higher than expected revenue without any increase in effort. Note, however, that the representative could alternatively reduce their timeliness rather than increase effort. If the representative relaxes their concern for timeliness, increasing wait time and expected past due benefit to \$5,300, the expected revenue is $0.6 * 5,300 = 3,180$. The return to increasing ρ is \$530 under the new cap (\$3,710-\$3,180), which is only \$130 higher than the return to increasing ρ under the old cap. For cases that were at the pre-existing cap, the return to decreasing timeliness under the new cap is \$780. As such, the return to decreasing timeliness in response to the cap increase is approximately six to seven times larger than the return to increasing ρ. In order for the fee cap to change effort, the effort cost to the representative must be more than \$400 but less than \$530. We will look for supporting evidence of each of these channels directly in the data.

These representative incentives have the potential to impact claimant welfare. Longer wait times could result in claimants spending longer periods without income, potentially leading to even worse health. Even though benefits accrued during this waiting period would be eventually paid out in a lump sum as past due benefits, this might reduce welfare if claimants face binding borrowing constraints. Longer wait times that result in more past due benefits and more dollars of representative compensation would also result in claimants paying a larger share of lifetime benefits to representatives. These incentives also affect claimants in the income distribution differently. Past due benefits are increasing in Average Indexed Monthly Earnings (AIME), meaning the incentive 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.

4 Data

We use the 2015 and 2018 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.[11] 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.[12] 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 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. To maintain the anonymity

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

[12] We combine the random samples from both the 2015 and 2018 files.

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).

Unfortunately, better data are not available to explore these relationships. Even the detailed administrative data Hoynes et al. (2021) use, only starts linking applicants to representative information starting in 2010, after the last maximum representative fee increase. Even with these data limitations, we provide the first estimates of how changes in the representative fee structure affects wait times for awardees.

We restrict the sample to only include SSDI recipients who applied in 1996 or later, so that we can accurately identify wait times. We also limit the sample to claimants whose application date was at least five years before the data collection date (either 2015 or 2018) to focus on resolved cases that are not truncated. We also focus on workers and exclude spouses and dependents (who cannot be linked to workers).[13] We make several other restrictions to

[13]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

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.[14]

5 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. 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.[15] We will first explore how changes in the maximum fee threshold affect wait times for claimants that are more likely and less likely to face a binding fee threshold.

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

multiple dependents).

[14]Data are also collected for SSI recipients. However, the DAF PUF does not include sufficient information on SSI application, onset, and entitlement dates to construct accurate measures of wait time. We cannot examine how representative compensation affects outcomes of SSI only claimants.

[15]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.

the current maximum fee threshold by solving the following equation

$$Monthly\ Benefit_i * Months * 0.25 = Threshold \tag{2}$$

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 the maximum fee threshold is a dollar amount, it 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. There are far fewer claimants at these high benefit levels. 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 the first panel of 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 the second panel of Figure 3, 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 (e.g. the coefficient on the quadratic term is larger) than the first panel of 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). Table A1 formally estimates the bindingness of the increases for all affected versus unaffected claimants and by primary diagnosis code.

5.1 Strategy 1: Difference-in-Differences

Figures 3 motivates 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:

$$Outcome_{it} = \alpha_0 + \beta(Affected_i * PostIncrease_t) + \gamma Affected_i + \phi PostIncrease_t + X_i\Gamma + \varepsilon_{it} \quad (3)$$

When estimating the effect of the 2002 increase, we limit the sample to claimants who applied between 1996 and 2009 (plausibly under either the $4,000 or $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 between 2002 and 2013 (plausibly under the $5,300 or $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 calculated using the claimant's past due benefits 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 Figure 3 shows 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:

$$Outcome_{it} = \alpha_0 + \phi_t + \sum_{\tau=1997}^{2013} \beta_\tau Affected_i * \phi_\tau + X_i\Gamma + \varepsilon_{it} \quad (4)$$

$Outcome_{it}$, $Affected_i$, 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 β_τ coefficients 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.

5.2 Strategy 2: Bunching

We will also explore how claimant outcomes change in 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 Appendix Figure A3). This kink point will vary across individuals, by benefit level.[16] 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

$$Month\ Waited\ Past\ Kink_i = Months\ Entitlement\ to\ Allowance_i - \frac{Threshold_{yr} * 4}{Benefit\ Level_i} \quad (5)$$

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 A4 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

[16] As seen in the second panel of Figure A3, the distribution of wait times is compressed towards zero for higher benefit claimants.

applicable Fee Threshold in equation (5). As seen in Figure 7, 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

$$Percent\ of\ Claimants_{mt} = \sum_{\tau=-60}^{60} \beta_\tau(\tau\ months\ from\ kink_m) * PostIncrease_t + \delta PostIncrease_t + \phi_m + X_{mt}\Gamma + \varepsilon_{mt} \quad (6)$$

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 claimant 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 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.

6 Results

6.1 Difference in Differences Results

Table 1 displays the estimated effects from equation (3). Column 1 includes controls for claimant demographics but excludes primary diagnosis fixed effects, column 2 includes primary diagnosis fixed effects, column 3 includes all controls but excludes claimants with benefit levels above $1,500 per month, and column 4 includes all controls but excludes claimants with benefit levels above $1,000 per month.[17] 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 include diagnosis fixed effects in column 2. 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

[17] Table A2 displays results excluding demographic controls and primary diagnosis fixed effects.

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 including diagnosis fixed effects 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 4 and 5 display the event study coefficients for these two outcomes. Figure 4 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 4 and 5 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 (3), 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 five groups (less than high school, high school, some college, a college degree, and education missing) while there are 10 regions that roughly correspond to census regions. As seen in Table A3, including labor market controls when looking at wait times yields similar point estimates, again implying an additional 0.4-0.7 months of wait time. In Figure 6, we plot region-specific treatment effects against a measure of region-level recession intensity, the peak-to-trough change in the employment to population ratio. We do not find a relationship between how hard a region was hit by a recession and our estimated effect of the fee increase in that region. In fact, the direction of the relationship suggests that we find smaller estimates in regions with stronger downturns, suggesting any confounding from these recessions will bias our estimates downward.

Finally, Table A1 shows how binding the pre-increase caps are for the four major diagnosis groups: mood disorder, internal systems, musculoskeletal, and infection. We find that among the four groups, claimants with a diagnosis for a disability related to infection are the least likely to reach the pre-increase caps. This is because those cases are resolved more quickly than the cases for other diagnosis groups. As a result, claimants with benefit levels above \$500 are not substantially more likely to reach the pre-increase cap than claimants with benefit levels below \$500. We view this group as providing a useful placebo test for our empirical design. In Table 3, we show heterogeneity in our results by primary diagnosis. We find increases in wait times for all major diagnosis groups except those claimants with a diagnosis for a disability related to infection. This lends further support to the validity of our difference-in-differences strategy and suggests the estimated increase in wait times is

not simply driven by differential trends by income.

We also examine whether the increase in wait times could be due to representatives putting more effort into cases. This does not appear to be the case. As seen in Appendix Figure A1, there are not discrete changes in successful claimant characteristics after fee ceiling changes as we might expect if representatives increase effort and change the probability of disability award. Instead, average characteristics seem to trend over the entire sample period. As an additional marker of effort, we explore the probability that claimants have an early onset date (prior to the application date). After the fee ceiling increase, representatives could increase past due benefits and fee payouts by providing more evidence that the disability began earlier.[18] However, as seen in Appendix Figure A2, the propensity of early onset dates does not seem to respond to the policy change, but trends in the opposite direction.[19] Although not conclusive, we do not find evidence that representatives are increasing wait times by increasing effort.

6.2 Distributional Bunching Results

We plot the estimated wait time density change in 2002 relative to 2000-2001 with 95-percent confidence intervals in Figure 8. We see a significant increase in density in the 12 months before and after the Months Wait Past Kink threshold. In some month bins, we observe increases in the number of claimants of up to 0.2 percentage points. 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

[18] It is worth noting this would be beneficial to the claimant as well as the representative.

[19] This is consistent with the early onset rate of low-benefit claimants catching up to early onset date of high-benefit claimants.

pattern is consistent with the increase in maximum representative compensation dragging out wait times.

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 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 style 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

$$\textit{Percent of Claimants}_{mt} = \sum_{\tau=1997}^{2008} \beta_\tau(-12 \leq m \leq 12) * (t = \tau) + \delta(-12 \leq m \leq 12) + \alpha_t + X_{mt}\Gamma + \varepsilon_{mt} \quad (7)$$

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

$$\textit{Percent of Claimants}_{mt} = \sum_{\tau=2003}^{2013} \beta_\tau(-18 \leq m \leq 0) * (t = \tau) + \delta(-18 \leq m \leq 0) + \alpha_t + X_{mt}\Gamma + \varepsilon_{mt} \quad (8)$$

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 Appendix Figure A5. In both cases the pre-trends are flat, but begin to turn upward and become significant one to two years before the policy change. 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. 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.

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. Consistent with the difference-in-differences results, if we control for the average employment to population ratio, and regional and education composition in each one month bin, the estimates are nearly identical (see Appendix Figure A5), suggesting this is not driven by changes in economic conditions.

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 creates 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 use difference-in-differences and bunching strategies to provide suggestive evidence

that the second feature of representative compensation rules leads to longer wait times for claimants. We find that the maximum representative fee increase in 2002 led to a 0.7 month increase in wait times while the 2009 increase led to an additional 0.4 months.

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, 15 percent of claimants retained representation (Hoynes et al., 2021). 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 4.9 month increase in wait times for claimants with representation while the 2009 increase was associated with a 2.9 month increase (assuming 15 percent take-up). Although information is limited, it is likely take-up at later stages (e.g., Reconsideration or appeal) are higher, suggesting this is an upper bound on the monthly increase.

Data limitations in the DAF PUF preclude us from answering interesting questions that could be examined using richer data. Data on individual claimant representation would lend itself to alternative identification strategies. Detailed information on representation agreement entry date could also be used to examine high frequency variation in close proximity to the maximum fee threshold changes. Information on the actual representatives could also reveal whether wait time effects are concentrated among certain representative types or groups or are broadly observed. Richer data could further shed light on 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 representatives delaying or withholding documentation to prolong the process. As noted above, much of the administrative data that contains representative information does not go back far enough to examine the effect of maximum representative fee increases (Hoynes et al., 2021).

This work sheds new light on how the incentives of social program gatekeepers can affect applicant outcomes. Many disability insurance claimants rely on representatives' expertise to successfully navigate and participate in the program. Current SSA regulation of representative compensation creates incentives to delay and results in longer wait times for claimants. These results also apply to other social programs and legal representative settings. Many programs rely on administrators and intermediaries, and policy architects should consider their incentives and objectives when designing social programs.

References

Agan, A., Freedman, M. and Owens, E. (2021), 'Is your lawyer a lemon? incentives and selection in the public provision of criminal defense', *The Review of Economics and Statistics* **103**(2), 294–309.

Autor, D., Kostol, A., Mogstad, M. and Setzler, B. (2019), 'Disability benefits, consumption insurance, and household labor supply', *American Economic Review* **109**(7), 2613–2654.

Autor, D., Maestas, N., Mullen, K. and Strand, A. (2017), 'Does delay cause decay? the effect of administrative decision time on the labor force participation rate and earnings of disability applicants', *NBER Working Paper Series No. 20840* .

Coe, N. B., Lindner, S., Wong, K. and Wu, A. Y. (2014), 'How do people with disabilities cope while waiting for disability insurance benefits?', *IZA Journal of Labor Policy* **3**(1).

Deshpande, M., Gross, T. and Su, Y. (2020), 'Disability and distress: The effect of disability programs on financial outcomes', *American Economic Journal: Applied Economics* **forthcoming**.

Deshpande, M. and Li, Y. (2019), 'Who is screened out? application costs and the targeting of disability programs', *American Economic Journal: Economic Policy* **11**(4), 213–248.

Federal Register (2009), 'Federal register social security administration'. Docket No. SSA-2009-0007.

Foote, A., Grosz, M. and Rennane, S. (2018), 'The effect of lower transaction costs on social security disability insurance application rates and participation', *Journal of Policy Analysis and Management* **38**(1), 99–123.

French, E. and Song, J. (2014), 'The effect of disability insurance receipt on labor supply', *American Economics Journal: Applied Economics* **6**(2), 291–337.

Gelber, A., Moore, T. and Strand, A. (2017), 'The effect of disability insurance payments on beneficiaries's earnings', *American Economic Journal: Applied Economics* **9**(3), 229–261.

Herd, P. and Moynihan, D. (2020), 'How administrative burdens can harm health', *Health Affairs* .

Hoynes, H., Maestas, N. and Strand, A. (2021), 'Legal representation in disability claims', *NBER Disability Research Center Paper* .

Lee, A. (2021), 'Flat fee compensation, lawyer incentives, and case outcomes in indigent criminal defense', *Working Paper* .

Maestas, N., Mullen, K. and Strand, A. (2013), 'Does disability insurance receipt discourage work? using examiner assignment of estimate causal effects of ssdi receipt', *American Economic Review* **103**(5), 1797–1829.

Maestas, N., Mullen, K. and Strand, A. (2015), 'Disability insurance and the great recession', *American Economic Review Papers and Proceedings* **105**(5), 177–182.

Michaud, A., Moore, T. J. and Wiczer, D. (2019), 'The relationship between social security disability insurance wait times and applications', *NBER Working Paper Series No. DRC NB18-Q5* .

Social Security Administration (2019), 'Annual statistical report on the social security disability insurance program, 2018'. Report A-12-14-24092.

Social Security Administration (2020), 'Annual statistical report on the social security disability insurance program, 2019'. Report A-12-14-24092.

Social Security Advisory Board (2012), 'Filing for social security disability benefits: What impact does professional representation have on the process at the initial application level?'.

Tables and Figures

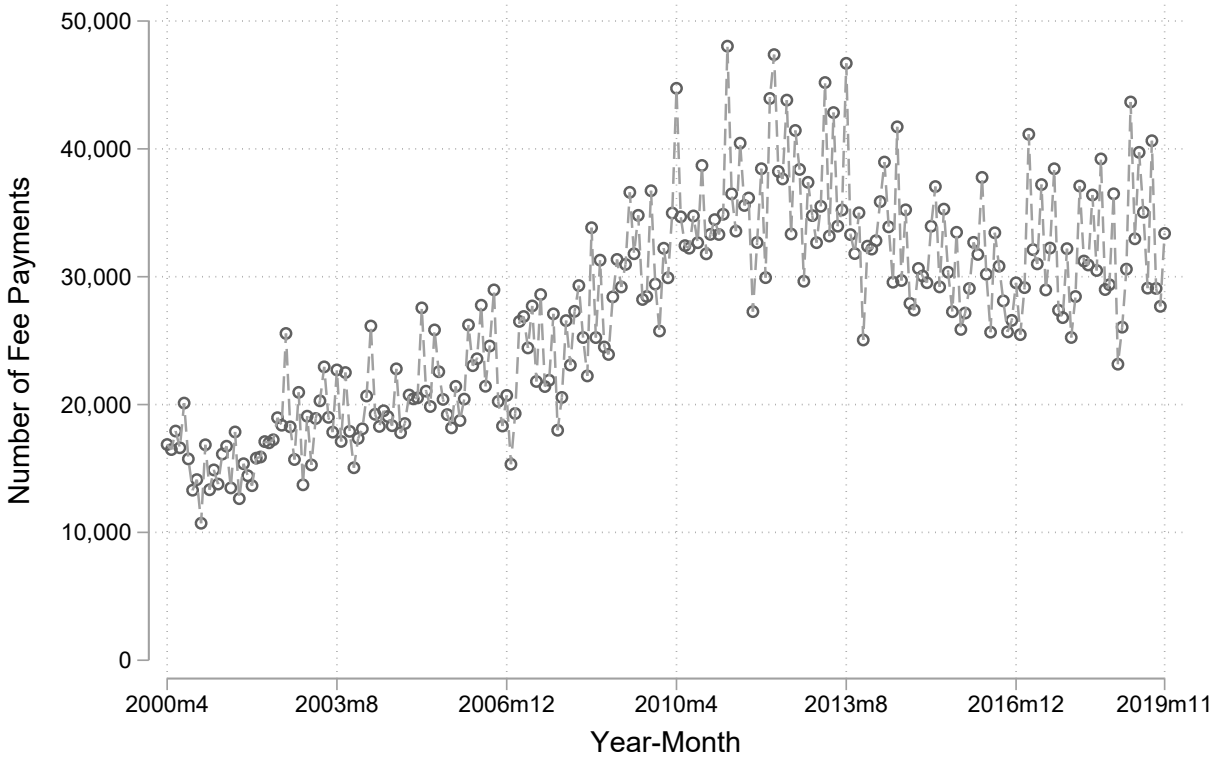

Figure 1: Number of Payments Made to Claimant Representatives by Month

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"

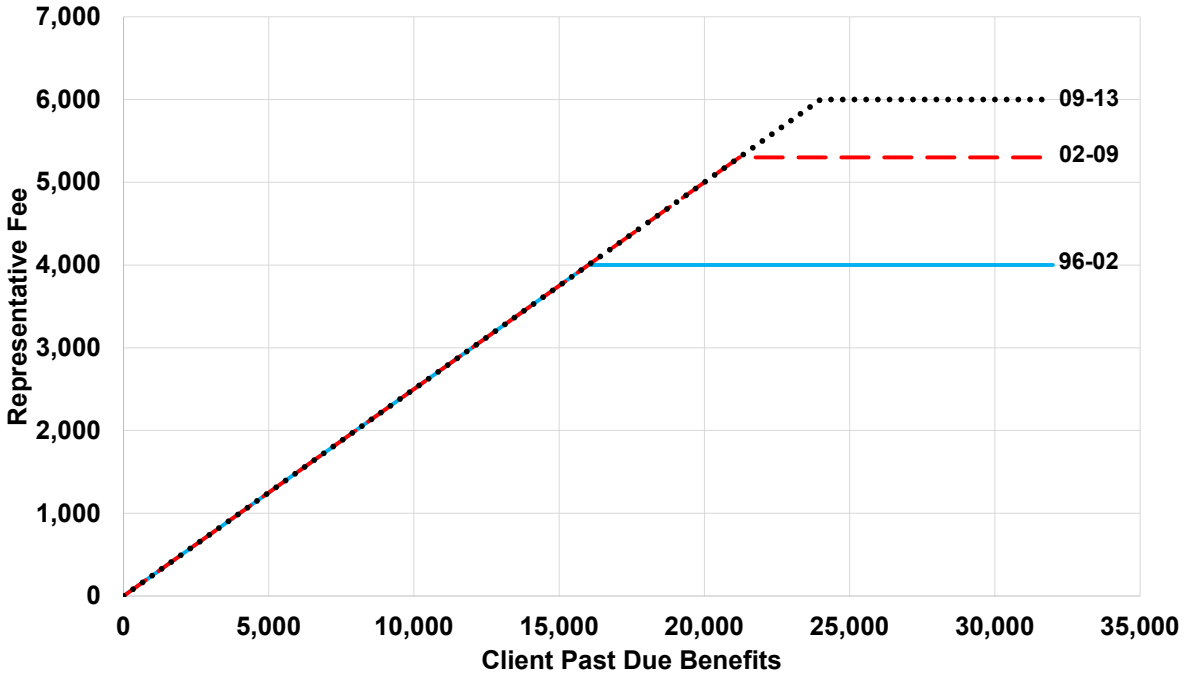

Figure 2: Representative Fee Schedule Over Time

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.

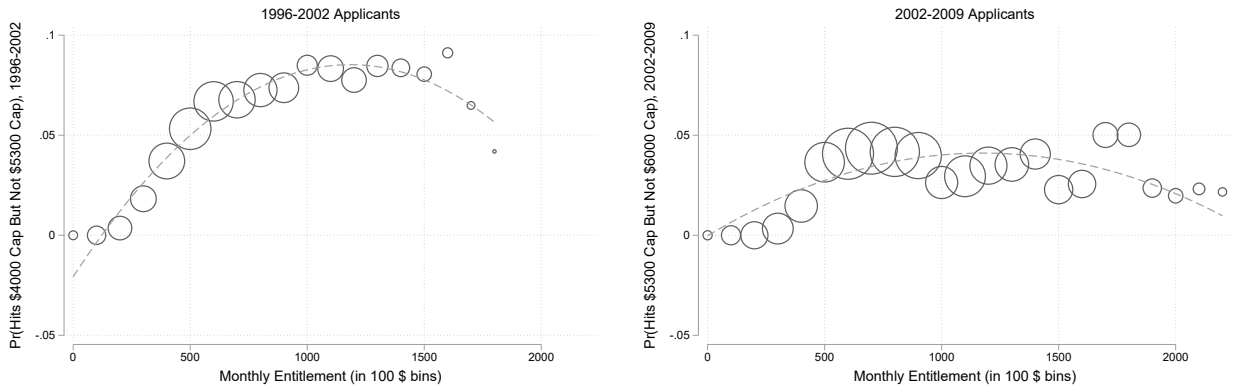

Figure 3: Share of Claimants from Pre-period Who Reached the Existing Fee Ceiling but Would Not Have Reached the New Fee Ceiling

Notes: Within 100 dollar monthly benefit bins we calculate the share of claimants for whom 25 percent of past due benefits exceeds the prevailing fee ceiling but would not have exceeded the new fee ceiling. In the left panel we include claimants from January 1, 1996 to February 1, 2002 who reached the $4,000 fee ceiling but would not have reached the $5,300 fee ceiling. In the right panel we include claimants from February 1, 2002 to June 22, 2009 who reached the $5,300 fee ceiling but would not have reached 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. 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.

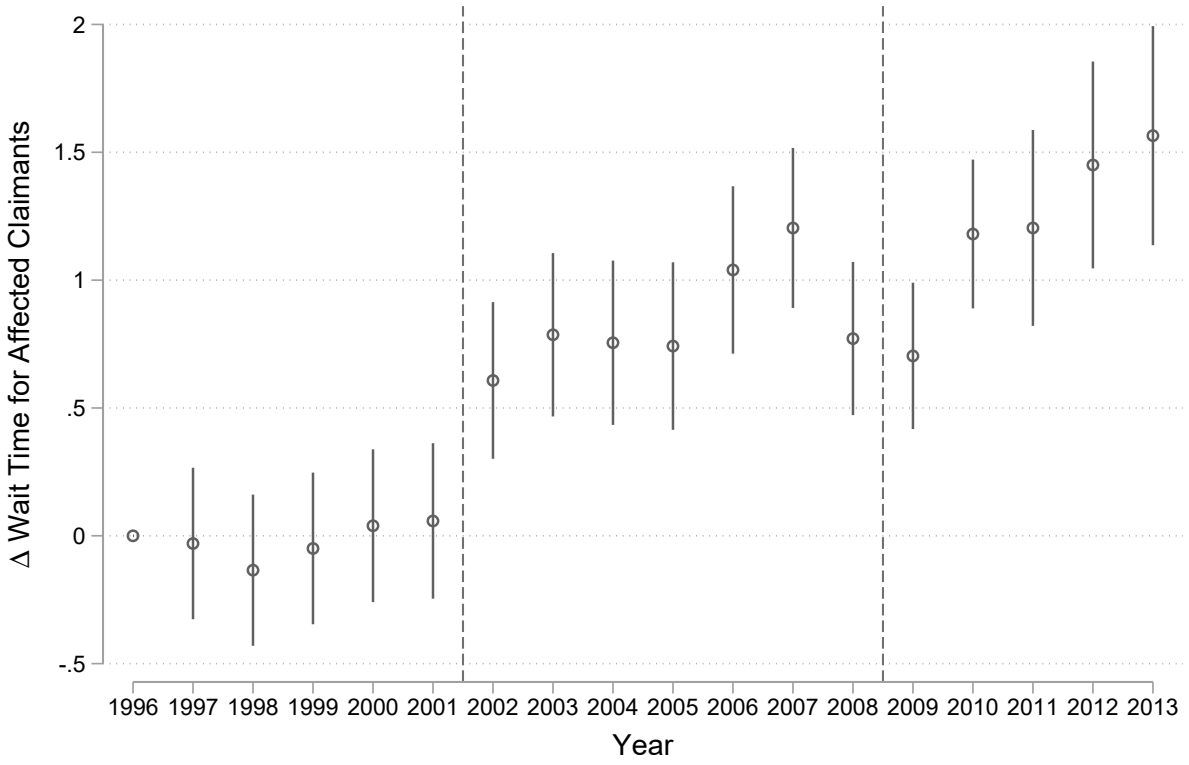

Figure 4: Change in Wait Time (in Months) For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). 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.

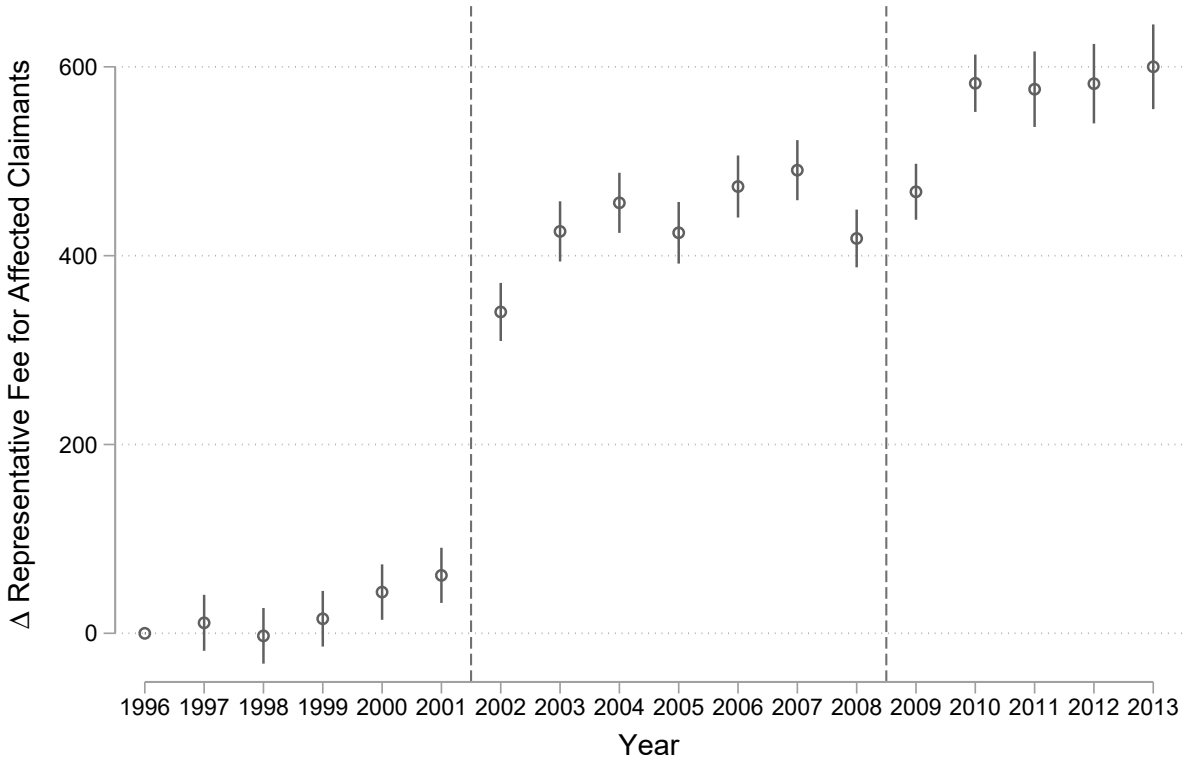

Figure 5: Change in Representative Fee (in Dollars) For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). 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.

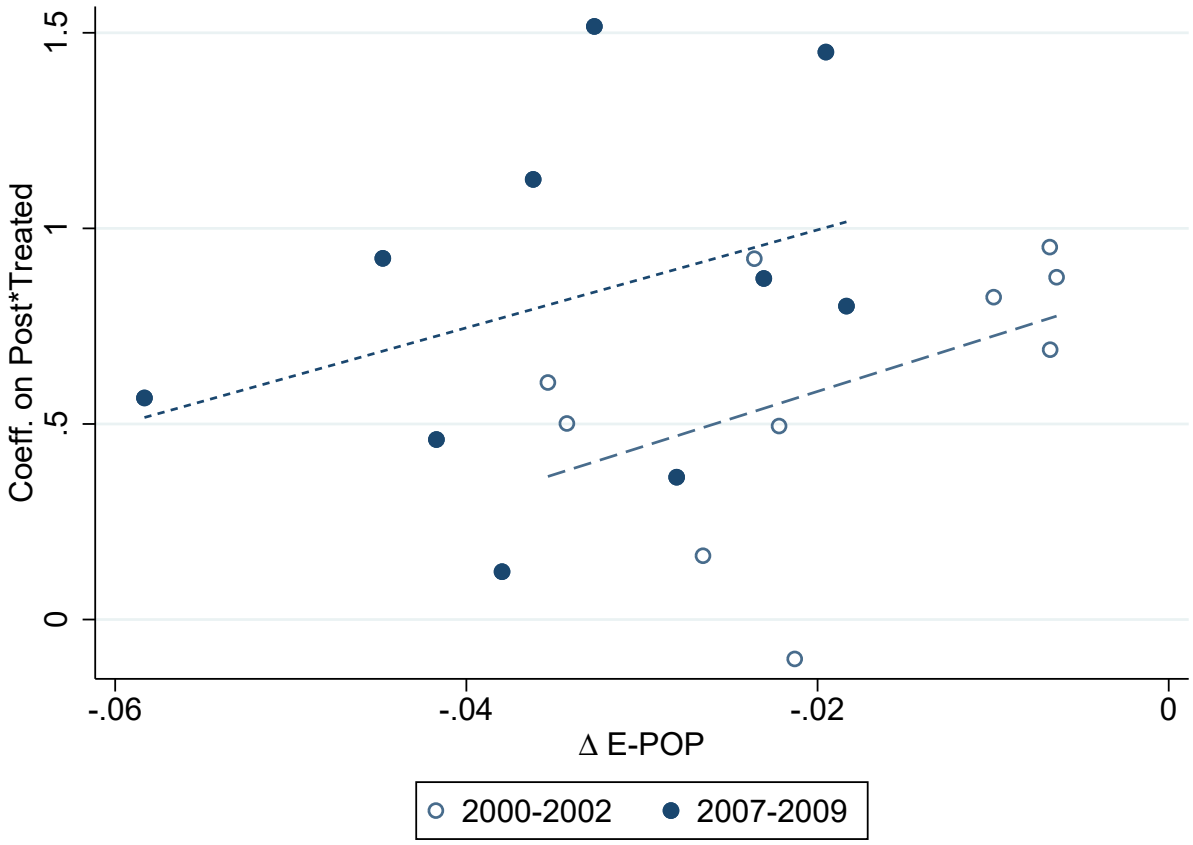

Figure 6: Relationship between Region Estimates and Economic Conditions

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated by region from equation (3). Each coefficient represents a region-level estimate of the change in wait time after the fee cap increase 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, and $100 monthly benefit bin fixed effects. These coefficients are plotted against the region-level change in the employment to population ratio from 2000-2002 and 2007-2009. The long dashed line plots the linear fit between the region-level estimates based on the 2002 fee increase and the 2000-2002 change in the employment to population ratio. The short dashed line plots the linear fit between the region-level estimates based on the 2009 fee increase and the 2007-2009 change in the employment to population ratio.

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

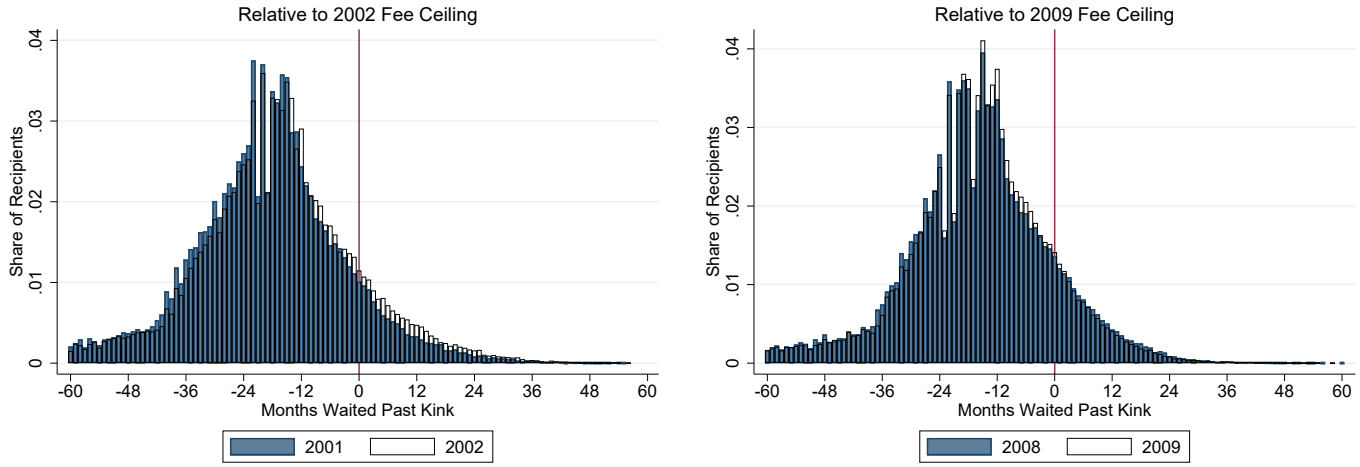

Figure 7: Distribution of Entitlement to Award "Wait Time" Relative to Maximum Representative Fee Threshold for Claimants Just Before and After Fee Ceiling Increase

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. In the left panel claimants that applied in 2001 and 2002 are plotted, relative to the new, 2002 fee schedule. In the right panel, claimants that applied in 2008 and 2009 are plotted, relative to the new, 2009 fee schedule. 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.

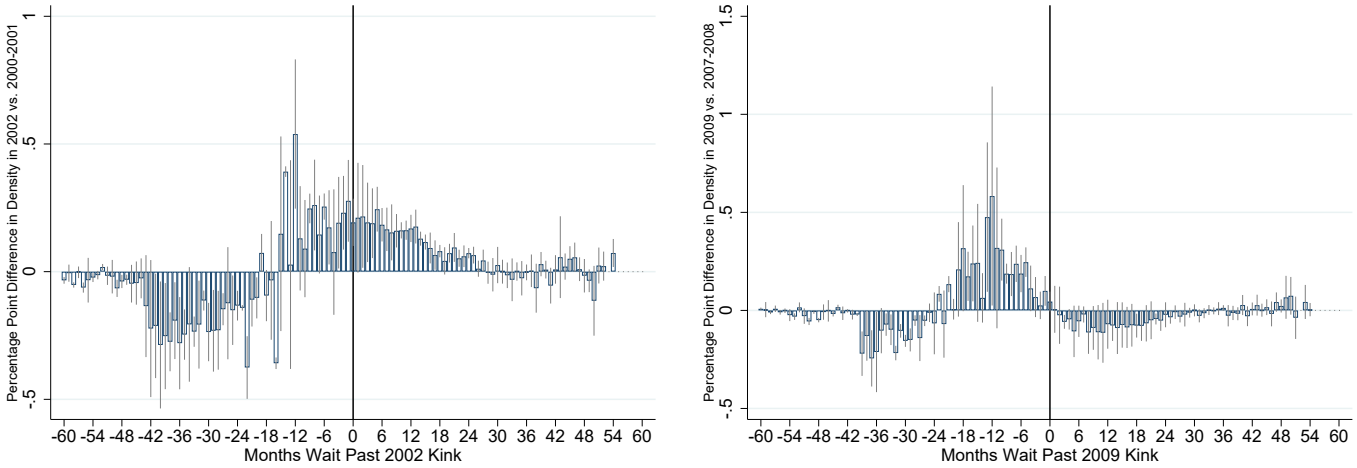

Figure 8: Difference in "Wait Time" Density Around New Fee Threshold in Pre- and Post Years

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 in the left panel. Only claimants that applied in 2007, 2008, or 2009 are included in the right panel. Each bar is the interaction coefficient from equation (5) with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial benefit level, and primary diagnostic code shares.

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

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	0.967***	0.735***	0.767***	1.131***
	(0.0657)	(0.0629)	(0.0631)	(0.0656)
Constant	16.41***	8.458***	8.188***	7.930***
	(0.134)	(0.444)	(0.467)	(0.710)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.053	0.145	0.141	0.127
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	0.639***	0.440***	0.373***	0.236***
	(0.0802)	(0.0782)	(0.0791)	(0.0825)
Constant	19.45***	12.20***	11.77***	11.36***
	(0.188)	(0.522)	(0.573)	(1.012)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.046	0.123	0.112	0.095
Primary Diagnosis FEs	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 (3) 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 (3) 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 includes controls for claimant sex and age and column 2 includes those controls plus primary diagnosis fixed effects. Column 3 includes the same controls as column 2 but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes those same controls 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	454.5***	417.3***	419.1***	425.8***
	(6.484)	(6.259)	(6.275)	(6.584)
Constant	2,502***	1,009***	979.8***	1,045***
	(15.42)	(107.4)	(107.5)	(139.2)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.090	0.192	0.193	0.191
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	175.9***	137.3***	113.3***	60.98***
	(8.299)	(8.266)	(8.401)	(8.941)
Constant	3,269***	1,670***	1,558***	1,524***
	(22.76)	(122.1)	(126.2)	(199.8)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.055	0.143	0.143	0.144
Primary Diagnosis FEs	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 (3) 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 (3) 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 includes controls for claimant sex and age and column 2 includes those controls plus primary diagnosis fixed effects. Column 3 includes the same controls as column 2 but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes those same 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 by Primary Diagnosis

	Wait Time from Entitlement to Approval (in Months)			
Primary Diagnosis	Mood Disorder	Int. Systems	Musc. Skel.	Infection
	(1)	(2)	(3)	(4)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009				
Affected x Post-2002	0.781***	0.383***	0.512***	-0.287*
	(0.115)	(0.126)	(0.149)	(0.151)
Constant	11.72***	13.95***	33.26***	9.385***
	(0.220)	(0.258)	(0.334)	(0.286)
Observations	284,946	339,420	256,576	188,140
R-squared	0.021	0.051	0.076	0.052
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	0.246	0.809***	0.447***	0.204
	(0.153)	(0.162)	(0.151)	(0.227)
Constant	14.57***	16.77***	39.17***	11.08***
	(0.314)	(0.371)	(0.389)	(0.450)
Observations	269,792	338,962	321,975	168,692
R-squared	0.017	0.044	0.116	0.047

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 (3) 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 (3) 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, all columns include demograhic controls (sex and age). Columns 1-4 are limited to claimants with diagnoses of: mood disorder, int. systems, musculoskeletal, and infection, respectively. In this table, the outcome variable is the claimant's wait time from entitlement to claim approval (in months).

Appendix A. Supplementary Analyses (For Online Publication)

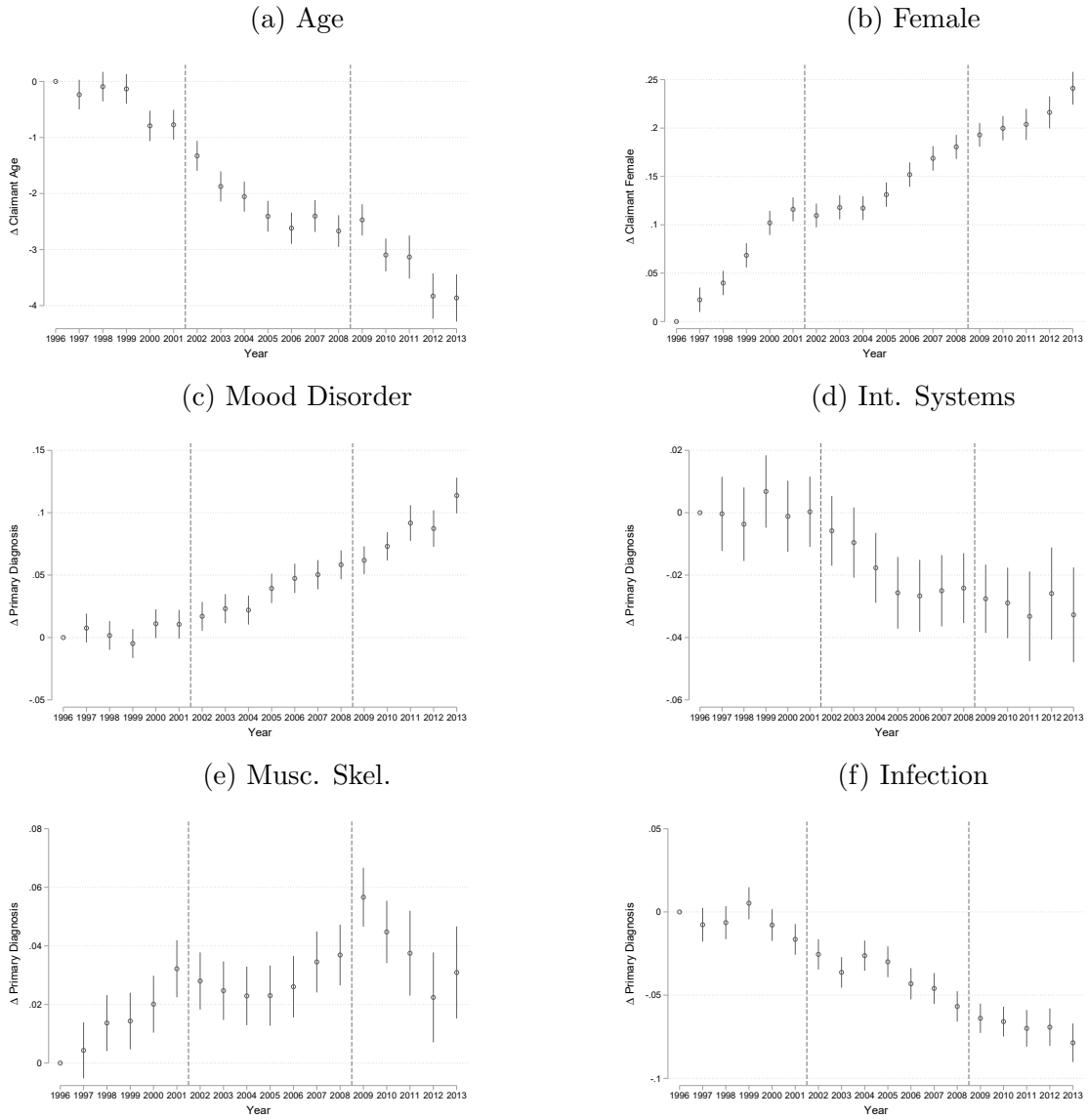

Figure A1: Claimant Characteristics for High Benefit vs. Low Benefit Claimants since 1996

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and applied in 1996 or later. The coefficients represent the change in the share of affected claimants with the given characteristic relative to unaffected claimants since 1996. Year of application fixed effects and $100 monthly benefit bin fixed effects are included.

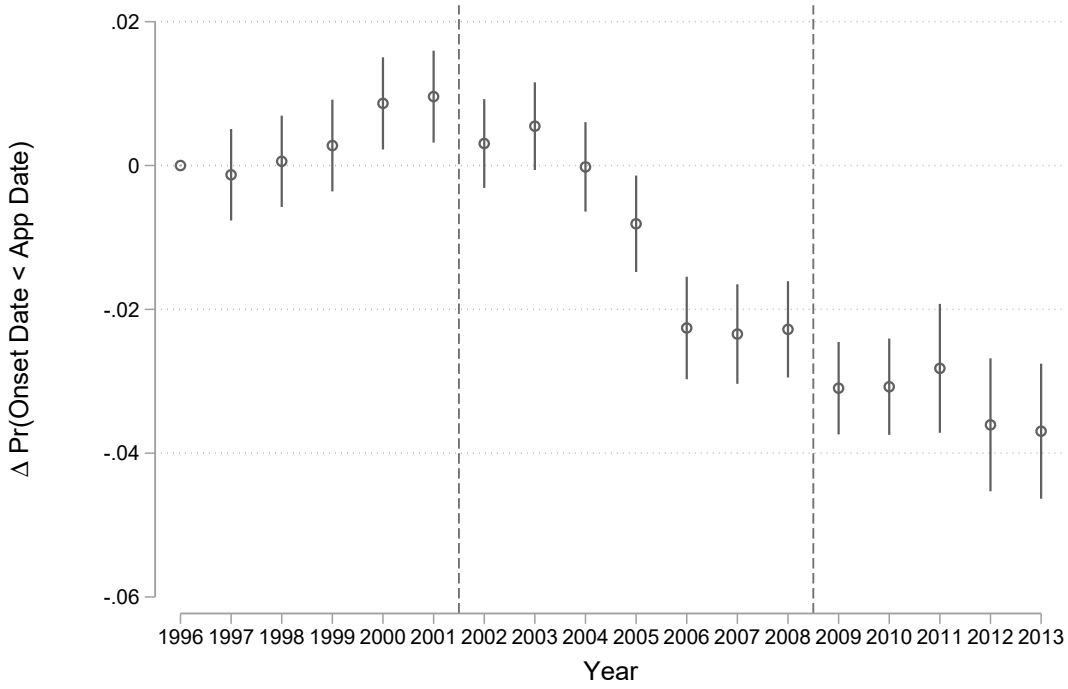

Figure A2: Change in Having Onset Date Before Application Date For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). Each coefficient represents the change in likelihood of having an onset date prior to application date 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.

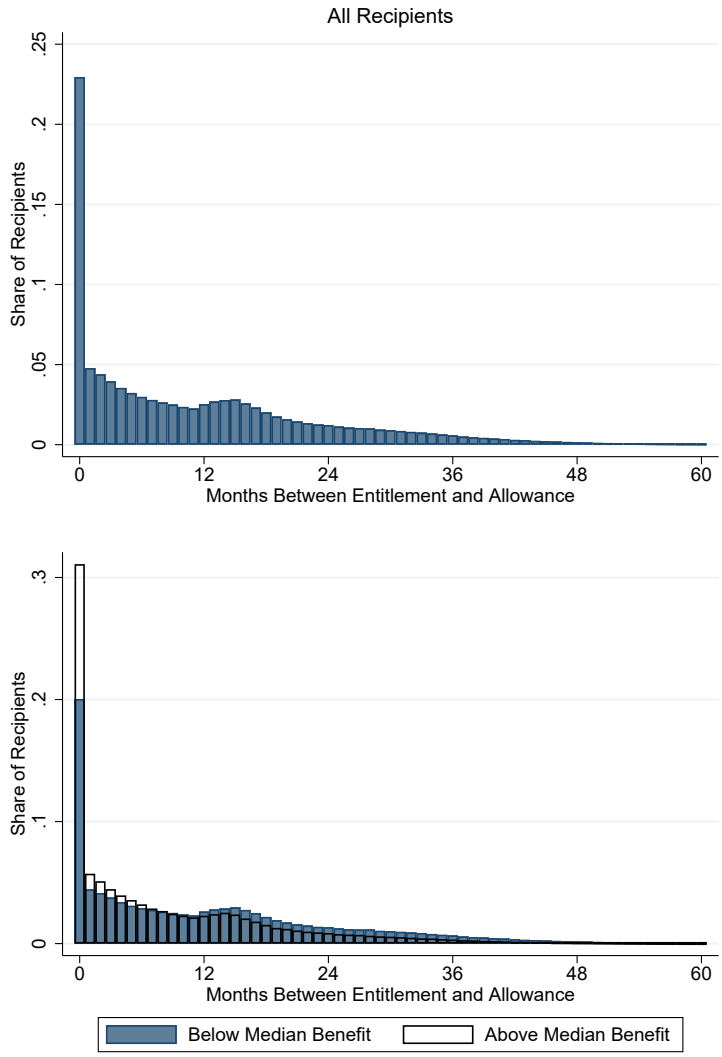

Figure A3: Distribution of Entitlement to Allowance "Wait Time" for Claimants

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.

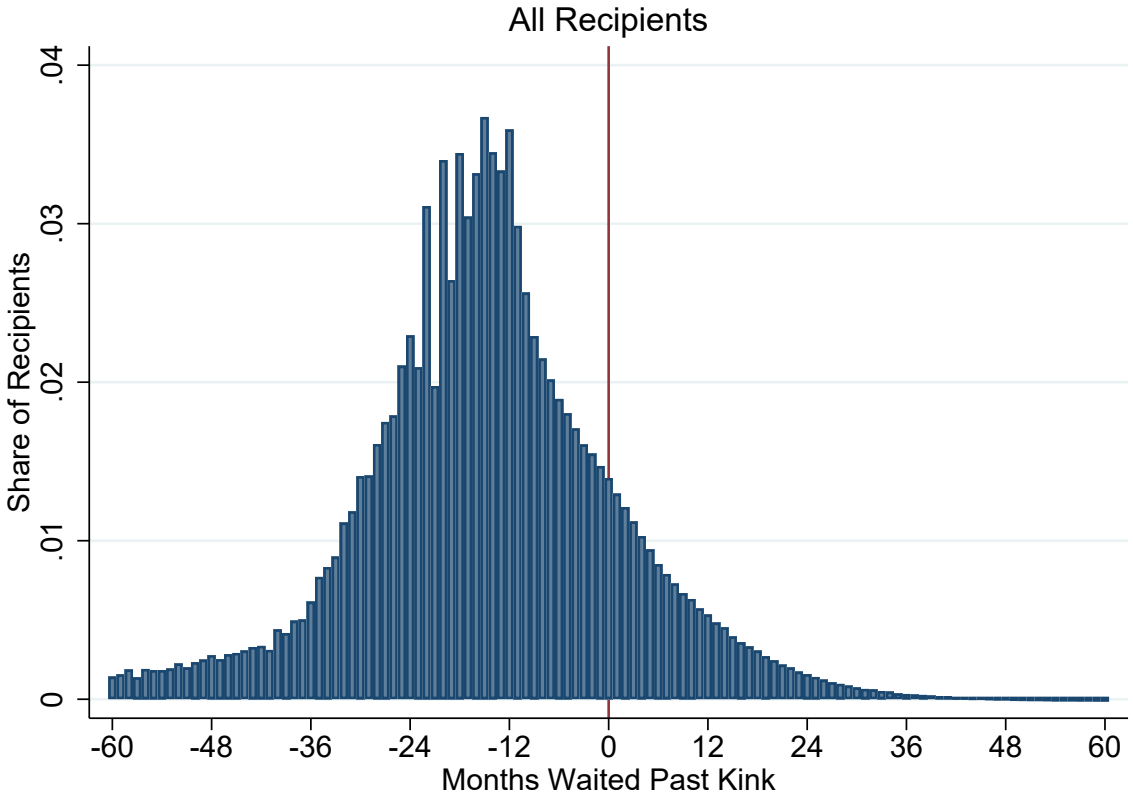

Figure A4: Distribution of Entitlement to Award "Wait Time" Relative to Prevailing Maximum Representative Fee Threshold

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date, applied between 1996 and 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 application year.

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

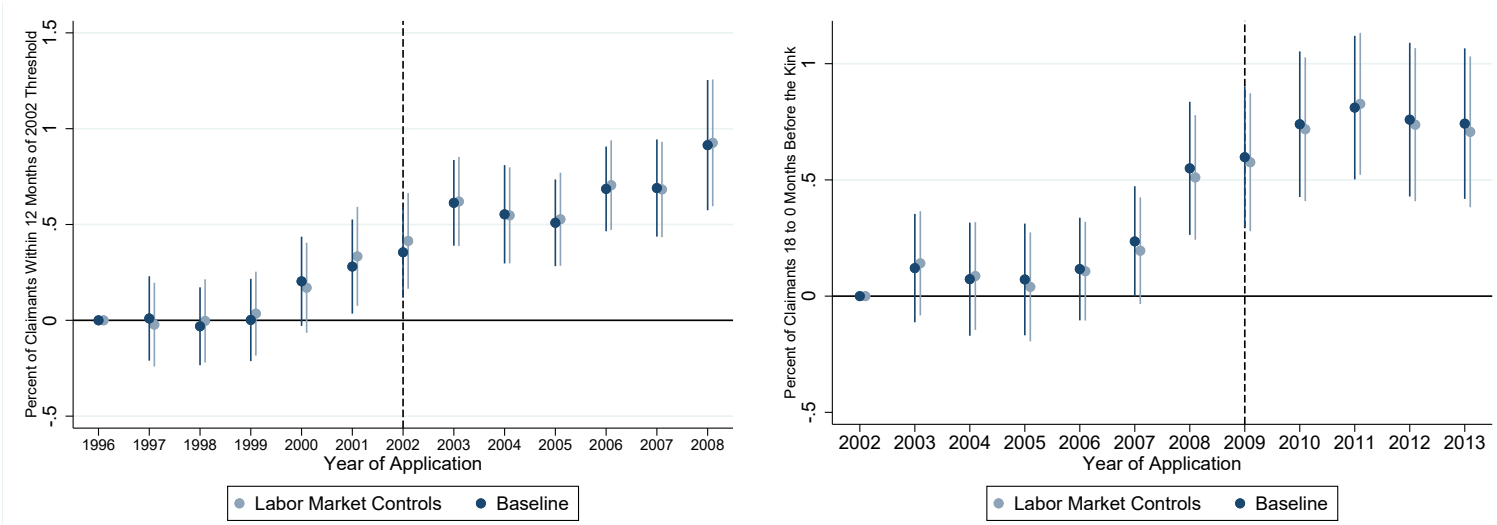

Figure A5: Event Study Difference in Bunching Around Maximum Fee Thresholds

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and 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 diagnostic code shares. The lighter plots also include education fixed effects, region fixed effects, and the average gender by education by region employment to population ratio in the bin and the share in each education and region bin to control for local labor market effects. Because the representative agreement date does not correspond to the application date, it is possible individuals who applied in the year leading up to the change face the higher maximum fee ceiling.

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

Table A1. Bindingness of Pre-Increase Fee Cap for Affected versus Unaffected Claimants

	Likelihood of Reaching Pre-Increase Cap but Not Post-Increase Cap				
	(1)	(2)	(3)	(4)	(5)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009					
Affected	0.0482***	0.0548***	0.0397***	0.0976***	0.0174***
	(0.000681)	(0.00134)	(0.00115)	(0.00215)	(0.00107)
Constant	0.0225***	0.0219***	0.0183***	0.0411***	0.00925***
	(0.000482)	(0.000843)	(0.000863)	(0.00151)	(0.000837)
Observations	379,237	95,724	119,947	67,963	72,942
R-squared	0.008	0.012	0.005	0.018	0.002
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013					
Affected	0.0306***	0.0296***	0.0266***	0.0471***	0.0137***
	(0.000347)	(0.000636)	(0.000631)	(0.000875)	(0.000648)
Constant	0.00663***	0.00612***	0.00548***	0.0108***	0.00278***
	(0.000254)	(0.000428)	(0.000490)	(0.000658)	(0.000515)
Observations	745,060	189,222	219,473	188,613	115,198
R-squared	0.003	0.004	0.002	0.005	0.001
Primary Diagnosis	All	Mood Disorder	Int. Systems	Musc. Skel.	Infection

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. In both panels, we define Affected as equal to one if the claimant's monthly benefits exceed $500 and zero if not. Panel A estimates the likelihood that a claimant reaches the $4,000 cap but would not the $5,300 cap for the sample of claimants who applied under the $4,000 cap. Panel B estimates the likelihood that a claimant reaches the $5,300 cap but would not the $6,000 cap for the sample of claimants who applied under the $5,300 cap. For both panels: column 1 shows these results for all diagnosis types. Columns 2-5 are limited to claimants with diagnoses of: mood disorder, int. systems, musculoskeletal, and infection, respectively.

Table A2. Effect of Fee Ceiling Increases on Claimant Wait Time, No Demographic or Diagnosis Controls

	Wait Time from Entitlement to Approval (in Months)		
	(1)	(2)	(3)
Panel A. 2002 Fee Ceiling Increase from $4,000 to $5,300, 1996-2009			
Affected x Post-2002	0.325***	0.769***	1.145***
	(0.0652)	(0.0656)	(0.0682)
Constant	13.20***	13.20***	13.20***
	(0.0417)	(0.0417)	(0.0417)
Observations	1,124,297	1,036,503	786,427
R-squared	0.021	0.023	0.023
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013			
Affected x Post-2009	0.200**	0.242***	0.164*
	(0.0806)	(0.0815)	(0.0848)
Constant	15.84***	15.84***	15.84***
	(0.0426)	(0.0426)	(0.0426)
Observations	1,145,400	991,716	685,114
R-squared	0.013	0.010	0.007
Sample Restriction	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 (3) 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 (3) 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, all specifications **exclude** demographic controls (sex and age) and primary diagnosis fixed effects. Column 2 excludes claimants with benefit amounts above $1,500 per month, and column 3 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 A3. 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	0.735***	0.758***	0.785***	1.167***
	(0.0629)	(0.0640)	(0.0641)	(0.0666)
Constant	8.458***	6.639***	6.757***	7.391***
	(0.444)	(0.597)	(0.630)	(0.961)
Observations	1,124,297	1,099,143	1,012,850	767,929
R-squared	0.145	0.151	0.147	0.133
Panel B. 2009 Fee Ceiling Increase from $5,300 to $6,000, 2002-2013				
Affected x Post-2009	0.440***	0.424***	0.364***	0.247***
	(0.0782)	(0.0784)	(0.0793)	(0.0827)
Constant	12.20***	9.228***	9.086***	9.171***
	(0.522)	(0.668)	(0.727)	(1.265)
Observations	1,145,400	1,133,319	981,770	678,676
R-squared	0.123	0.130	0.120	0.103
Primary Diagnosis FEs	YES	YES	YES	YES
Labor Market 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 (3) 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 (3) 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 includes demographic controls (age and sex) and primary diagnosis fixed effects. Column 2 includes those controls as well as the annual gender by education by region employment to population ratio, region fixed effects, 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).