# The Economic Consequences of Bankruptcy Reform*

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**Abstract**

A more generous consumer bankruptcy system provides greater insurance against financial risks, but it may also raise the cost of credit to consumers. We study this trade-off using the 2005 Bankruptcy Abuse Prevention and Consumer Protection Act (BAPCPA), which raised the costs of filing for bankruptcy. We identity the effects of BAPCPA on borrowing costs by exploiting variation in the effects of the reform on bankruptcy risk across credit score segments. Using a combination of administrative records, credit reports, and proprietary market research data, we find that the reform reduced bankruptcy filings and that the reduction in filings led to a decrease in credit card interest rates, with an implied pass-through rate of about 50–75 percent. We interpret our results using a stylized model of consumer bankruptcy and conclude that the estimated pass-through effects are consistent with the marginal consumers deterred by BAPCPA repaying a large share of their unsecured debts instead of discharging their debts through bankruptcy.

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## 1 Introduction

In recent decades, the rate of consumer bankruptcy filings in the United States climbed from 0.3 percent of households annually in the early 1980s to 1.5 percent in the early 2000s (Board of Governors, 2006). This increase in bankruptcies was cited by lawmakers as a reason to pass the 2005 Bankruptcy Abuse Prevention and Consumer Protection Act (BAPCPA). The bill implemented a number of provisions that collectively made filing for bankruptcy more onerous, more expensive, and less financially beneficial.

By making bankruptcy less attractive, BAPCPA was widely expected to reduce bankruptcy filings. But there was considerable debate at the time regarding how the bankruptcy reform would affect consumer credit markets. Proponents of the bill argued that creditors would pass through higher debt-recovery rates in the form of lower interest rates. Judge Richard Posner argued that "the new Act…should reduce interest rates and thus make borrowers better off" (Posner, March 2005). In Senate testimony, Todd Zywicki argued that "We all pay for bankruptcy abuse in higher down payments, higher interest rates, and higher costs for goods and services" (House Report, 2005). By contrast, critics of the reform argued that any reduction in filings would not be passed-through to borrowers but would instead be captured by lenders. Additionally, critics contended that claims of abuse were overstated, and that the bill would harm filers struggling with medical expenses and job loss (Warren and Tyagi, 2005).

This paper informs that debate by estimating how the decline in bankruptcy filings caused by BAPCPA affected the interest rates paid by consumers in unsecured consumer credit markets. Our analysis consists of three main steps. First, we document that the reform dramatically decreased the number of bankruptcies. While the decline in bankruptcies is clearly evident in the raw time series data, we discipline our empirical analysis using "excess mass" techniques borrowed from the recent literature studying bunching at tax kinks (Saez 1999; Kleven and Waseem 2003). We do this in order to estimate the counterfactual bankruptcy filing trends in the absence of the bankruptcy reform and estimate the net decrease in filings over the longer run. In the months immediately before the reform, we estimate a net increase in filings of more than 750,000, as consumers rushed to file for bankruptcy to discharge their debts before the new bankruptcy code was implemented.

Over time, however, we find that the reform reduced the bankruptcy rate by roughly 50 percent; net of the short-run increase during the "rush to file," we estimate that there were roughly one million fewer bankruptcy filings in the two years after BAPCPA than would have occurred without the reform.

Second, we develop a stylized model of consumer bankruptcy to study the relationship between the risk of bankruptcy and the cost of borrowing. We use the model to predict how changes in the generosity of the bankruptcy regime (e.g., increasing the cost of filing or decreasing asset exemptions) should affect filing rates and borrowing costs. We show that when changes in the cost of filing affect bankruptcy rates, the pass-through to interest rates depends on the amount recovered from those deterred from filing at the margin. All else equal, the higher the average repayment rate for these marginal filers, the greater the change in interest rates.

Third, we use the model to motivate our empirical strategy which identifies the effect of BAPCPA on interest rates by exploiting variation in the effects of the reform on bankruptcy filings across credit score "segments." To do so, we first estimate the effect of BAPCPA on bankruptcy filings across the credit score distribution, using detailed credit report data that contains both credit scores and records of filing for bankruptcy. We then combine these credit-score-segment-specific estimates with proprietary data on credit card offers to estimate event-study and difference-in-difference regression models that compare credit-score segments before and after the reform. The key identifying assumption is that interest rates would have evolved similarly across credit score segments absent the reform. We provide evidence supporting this "common trends" assumption in the main event-study figure, which shows no differential trends across credit-score segments in the years leading up to the reform, and we also show that our results are robust to subprime-specific trends, suggesting that our results are not driven by other factors differentially affecting the subprime credit market during this time period.

We find that each one-percentage-point decline in the risk of filing for bankruptcy within a credit-score segment decreases average interest rates in that segment by 67 basis points. The estimated magnitude falls by about one-third when accounting for "teaser" interest rates (which are generally less variable across credit-score segments), but the estimates remain statistically and economically significant. Interpreting these estimates through the lens of our stylized model, we conclude that between 50 and 75 percent of the cost savings to creditors from the reduced bankruptcy filings were passed on to consumers. According to the model, these results imply that the marginal consumers deterred by BAPCPA likely repaid a large share of their unsecured debts instead of discharging their debts in bankruptcy.

One of the central provisions of BAPCPA was a "means test" which restricted the options available to high-income filers and was intended to "ensure that debtors repay creditors the maximum they can afford" (House Report, 2005). The means test was designed to shift filers with higher incomes from Chapter 7 to Chapter 13 by eliminating the option to file Chapter 7.[1] While we find that BAPCPA lowered interest rates, we find suggestive evidence that this effect was *not* in fact driven by a disproportionate decline in filing by high-income filers, as policymakers had intended. Instead (and counter to the intent of the law), we find no change in the distribution of income of bankruptcy filers when we proxy for income using the median income in the filer's ZIP Code.[2] These findings are consistent with other research which has suggested that the increasing cost of filing for bankruptcy deterred filings at least as much as the means test (Gross et al., 2014).

We also find suggestive evidence that BAPCPA reduced the rate of bankruptcy filing following a hospital admission. Building on the event-study approach of Dobkin et al. (2018a), we find that an uninsured hospitalization increases the likelihood of filing for bankruptcy by 1.5 percentage points prior to the reform, but by just 0.4 percentage points after the reform. While a hospitalization is just one example of an adverse health event (and the share of bankruptcies caused by hospitalizations appears to be small according to Dobkin et al. 2018b), to the degree that this finding generalizes to other types of adverse shocks, these results suggest that the bankruptcies deterred by BAPCPA were not limited to the most "abusive" filings.

Taken together, the results in this paper suggest that BAPCPA reduced bankruptcy filings, but not in the targeted way that proponents had intended. Nevertheless, the imperfect targeting of the law still appears to have provided meaningful benefits to many consumers in the form of lower credit

[1] Chapter 7 bankruptcy offers filers a "fresh start." All qualifying debts are discharged in exchange for their non-exempt assets. Chapter 13 bankruptcy offers filers a "reorganization." Chapter 13 filers lose no assets, but must commit to a repayment plan out of their future income.

[2] Additionally, we find only a small increase in the share of Chapter 13 filings and find that filings of both chapters declined after the reform. These results suggest that the BAPCPA caused large declines in bankruptcy rates across a broad range of consumers that swamped any effect of BAPCPA on inducing consumers to switch between Chapter 7 and Chapter 13.

card interest rates. Our model neatly reconciles these results by highlighting the importance of the average repayment rates of marginal filers in determining the change in interest rates. The model also reveals that high repayment rates for marginal filers is consistent with high costs of filing. This highlights the stark trade-off for policymakers: marginal filers may benefit substantially from a more generous bankruptcy system (getting a "fresh start" instead of repaying a larger share of their debts), but these benefits come at the cost of higher interest rates for other consumers.

This paper contributes to three main areas of research. First, it contributes new evidence on the effect of the bankruptcy code on consumer credit markets. Gropp et al. (1997), Berkowitz and White (2004), and Severino and Brown (2017) use cross-state variation in bankruptcy exemptions and find that more-generous exemptions are associated with less readily available credit, and Chakrabarti and Pattison (2016) find that the elimination of auto loan "cramdowns" under Chapter 13 reduced interest rates on auto loans.[3] Our paper is also related to earlier evaluations of BAPCPA which focused on aggregate interest-rate spreads (Simkovic, 2009) and student loans (Alexandrov and Jiménez, 2017). While our time-series analysis shows some apparent narrowing of interest rates between subprime and prime consumers, our main results go well beyond a simple time-series analysis by exploiting variation in effects of the reform across credit score segments. To our knowledge, this paper provides the first event-study estimates of the effects of BAPCPA on credit card interest rates.[4]

Second, this paper provides new evidence on pass-through in consumer credit markets. Existing research on pass-through in credit markets emphasizes sticky interest rates (and thus limited pass-through), but this research typically estimates pass-through using shocks to the cost of funds rather than shocks to default risk (e.g., Ausubel, 1991; Calem and Mester, 1995; Stavins, 1996; Stango, 2000; Calem et al., 2006; Agarwal et al., 2017). We thus contribute new evidence to this literature

[3] Cramdowns allowed borrowers to reduce the principal of their loan to the current market value of the vehicle.

[4] Other related work on BAPCPA includes studies by Li et al. (2011), Morgan et al. (2012), and Mitman (2016). Those papers argue that by reducing the substitutability between bankruptcy and foreclosure, BAPCPA increased foreclosures and exacerbated the mortgage crisis. Albanesi and Nosal (2018) also evaluate BAPCPA and document similar declines in bankruptcy filing rates for Chapter 7, which they attribute to liquidity constraints from the increased cost of filing. They also document an accompanying rise in insolvency and, along with Han and Li (2011), show that bankruptcy filers have better access to credit than the insolvent. Nakajima (2017) argues BAPCPA improved welfare by lowering default premia and improving consumption smoothing, but not for agents with temptation. Relative to this work, this paper contributes event-study estimates of the effect of BAPCPA on interest rates and estimates pass-through.

by estimating how changes in bankruptcy filing risk are passed-through to credit card interest rates. While a full reconciliation of the different pass-through estimates is outside the scope of this paper, Grodzicki (2017) argues that credit card markets have become more competitive in recent years following the costly adoption of screening technology, which is consistent with a large amount of price responsiveness to changes in default risk due to bankruptcy reform.

Lastly, this paper is related to studies of the consumer bankruptcy system that calibrate structural models (Zame, 1993; Dubey et al., 2005; Livshits et al., 2007; Chatterjee et al., 2007; Mitman, 2016; Nakajima, 2017). Those models typically emphasize the trade-off between using bankruptcy to smooth consumption across states of the world and the higher cost of smoothing consumption over time, often while assuming perfectly competitive credit markets (and thus full pass-through of the costs of lending to consumers). Our results suggest a large amount of pass-through from bankruptcy reform to interest rates in consumer credit markets, which provides empirical evidence in support of these modelling assumptions.

The remainder of this paper proceeds as follows. The next section provides background information on bankruptcy before and after BAPCPA. Section 3 develops a simple model to guide an assessment of the costs and benefits of bankruptcy reform. Section 4 describes our data sources and sample construction. Section 5 evaluates how BAPCPA affected the number of filings, Section 6 then estimates the pass-through of this decline in bankruptcy to borrowing costs. Section 7 evaluates how the reform changed *who* declares bankruptcy. Section 8 concludes.

## 2 Institutional Background

In contrast to other developed countries, American consumers have historically enjoyed an exceptionally debtor-friendly bankruptcy system.[5] In particular, American consumers filing for bankruptcy have had the option to freely choose between a "fresh start" (liquidating outstanding debts through Chapter 7) and a "reorganization" of debts (repaying debts on an installment plan over several years through Chapter 13). Chapter 7 filers must forfeit all non-exempt assets in exchange for discharge of their debt, while Chapter 13 filers are allowed to keep all of their assets

[5] Italy, for instance, had no form of consumer bankruptcy until 2015, and Germany only began allowing consumer bankruptcy in 1999. Before then, consumers in those countries had few options to discharge their debts (Tabb, 2005).

but must repay their debt out of future income.

Despite the potential financial benefits, consumer bankruptcy has historically been a relatively rare phenomenon in the United States. In the late 1970s, just 0.3 percent of households filed for bankruptcy in a given year. In 1978, the United States adopted a new bankruptcy code and legalized lawyer advertising. A 1978 Supreme Court decision allowed banks to export their home interest rates and so evade state usury laws.[6] These changes in policy catalyzed the growth of unsecured borrowing in the ensuing decades (White, 2007). By 1999, the bankruptcy rate had increased to 1.5 percent, prompting creditors to lobby for a more-stringent bankruptcy code.

To make their case, credit-industry lobbyists pointed to a handful of high-profile cases of "exemption shopping." For those bankruptcies, debtors moved across state lines to select the most-beneficial bankruptcy regime, and this was held up as emblematic of the abuse rampant in the bankruptcy system. A reform of the bankruptcy system was first drafted in 1998 and passed by Congress in 2000, but pocket-vetoed by President Clinton. The bill was reintroduced each Congress until it finally passed with broad bipartisan support in 2005. The Senate passed the bill on March 10, 2005, the House on April 14, 2005, and it was signed by President Bush on April 20, 2005. The new bankruptcy code went into effect for all bankruptcies filed on or after Monday, October 17, 2005.

BAPCPA made filing for bankruptcy less attractive in three primary ways. First, the law sought to prohibit higher-income households from filing Chapter 7. To do so, lawmakers introduced a means test which they referred to as "the heart of the bill" (House Report, 2005). The means test added a "presumption of abuse" for filers whose income is above certain thresholds. Debtors are subject to the means test if their income from the previous six full months before filing, adjusted for family size, is more than the state median income (Parra, 2018).[7] Debtors subject to the means test are functionally prohibited from filing Chapter 7, and can only file Chapter 13 (which, post

[6]In *Marquette National Bank of Minneapolis v. First of Omaha Service Corporation* (439 U.S. 299 (1978)), the U.S. Supreme Court ruled that state anti-usury laws regulating interest rates are not enforceable against nationally chartered banks based in other states.

[7]Virtually all income is included in this calculation with the notable exception of Social Security income. Those with debts that are not "primarily consumer debts" (e.g. business investments) are also exempt from the means test. Debtors can also "pass" the means test if they can demonstrate that their "disposable income" (income after allowed deductions) is less than $182.50 or $109.59, if that is enough to pay unsecured creditors more than 25 percent of the debt owed over five years.

BAPCPA, required higher repayment). This created an incentive for borrowers to suppress their labor supply and earnings below the state median in order to skirt the means test and file Chapter 7 or reduce their repayment obligation under Chapter 13.[8]

Second, BAPCPA limited the benefits of filing for bankruptcy along a number of dimensions. Prior to BAPCPA, Chapter 13 filers could propose their own repayment plan and faced no incentive to offer a repayment plan more generous than the relief they would receive under Chapter 7. After BAPCPA, Chapter 13 filers are required to forfeit 100 percent of their disposable income for five years to pay down their debts.[9] The reform also limited the ability of filers to discharge some purchases and "exemption shop" for the most-favorable state bankruptcy regime. Debtors who move must now wait two years before they are allowed to file under their new state's exemptions. Bankruptcy filers must wait a set number of years before they are allowed to file again. BAPCPA increased the waiting period from six years to eight years for Chapter 7 and from six months to two years for Chapter 13.

Finally, BAPCPA made the process of filing for bankruptcy much more burdensome and expensive. Bankruptcy court fees themselves increased. Bankruptcy filers are now required to take two educational courses: a credit-counseling course before filing and a financial-management course before their debt is discharged. Filing requirements increased and bankruptcy attorneys were made liable for any inaccuracies in the filing, which increased attorney costs by as much as $500 and subsequently increased the fees they charged their clients, which are not dischargeable in bankruptcy (House Report, 2005). Altogether, these changes increased the mean financial cost of filing from $868 to $1,309 for Chapter 7 and from $2,260 to $2,861 for Chapter 13 (Lupica, 2012).

[8]The incentive to suppress income prior to filing is relevant even if households cannot suppress it enough to get under the state median. Chapter 13 repayment plans, which are paid over the subsequent five years, are based on documented disposable income *over the prior six months*. As White (2007) points out, a reduction in monthly earnings of $1 for the six months prior to filing costs filers $6 in the short-run but reduces their repayment requirement by $60 ($1 each month over the next 60 months).

[9]Allowances for living expenses vary by metropolitan area and are largely based on the Internal Revenue Service policies for the treatment of delinquent taxpayers.

## 3 Economic Framework

We develop a simple economic framework to determine the effects of an increase in the cost of filing on bankruptcy filing rates and borrowing costs.[10] In addition, we use the framework to calibrate a benchmark for the expected magnitude of any potential effects on interest rates.

### 3.1 Model Set-up

There are $I$ different types of individuals, with individual types indexed by $i \in 1, ..., I$. The types are intended to represent the credit score "segments" that we use in our empirical analysis. There is a unit mass of each type, and individuals within each type are ex-ante identical and live for two periods. In the first period, each individual borrows $b_i$ at interest rate $r_i$, so that if the debt is repaid in full in the second period, then $(1 + r_i)b_i$ is repaid.[11] In the second period, each individual receives income $y$, drawn from distribution $f_i(y)$, with associated cumulative distribution function $F_i(y)$.

After observing their income, $y$, individuals can either repay their outstanding debt, or they can file for bankruptcy and retain all of their assets up to an exemption level, $e$. When individuals file, they must also pay cost $c$, which captures all relevant fixed costs of filing: filing fees, legal costs, hassle costs, and stigma.

This leads to a simple decision rule: an individual will file for bankruptcy if income (net of full repayment of debt) is less than the exemption amount minus the cost of filing $(e - c)$; that is, if $y - (1 + r_i)b_i < e - c$.[12] When individuals file for bankruptcy, creditors recover $\max\{0, y - e\}$. We work through the case where individuals with income below $c$ are insolvent and unable to file for bankruptcy in Appendix Section A.3, but for simplicity here, we assume that $f_i(y)$ has no support below $c$ for all types.

[10]The model we develop in this section is most closely related to the model in Wang and White (2000), which simulates various optimal bankruptcy reforms, but does not discuss the determinants of the magnitude of pass-through, which is one of the goals of our analysis.

[11]We treat $b_i$ as exogenous throughout this simple framework. In reality, changes in the bankruptcy code should also affect the amount borrowed. We make this simplification in order to focus on the response of creditors to the change in bankruptcy filings induced by BAPCPA. Given the relatively short-run nature of our empirical analysis, we view this as the first-order impact of the changes to the bankruptcy code for our purposes.

[12]Consistent with the fact that the costs of filing are not dischargeable, in the model $c$ must be paid out exempt assets.

### 3.2 Impact of Bankruptcy Reform on the Cost of Credit

Within the population of borrowers of each type, the probability of filing for bankruptcy can be defined as $p_i = F_i(e + (1 + r_i)b_i - c) = F_i(y_i^*)$, where $y_i^*$ represents the second-period income at which a borrower of type $i$ is indifferent between filing and not filing. Based on this decision rule, we can calculate how the exemption level $e$ and filing costs $c$ affect filing and repayment behavior. We are ultimately interested in the relationship between the bankruptcy code (i.e., $e$ and $c$) and the cost of borrowing ($r_i$), but we begin by estimating how changes in the exemption level and cost of filing determine the share of individuals who choose to file ($p_i$).

**Remark 1.** *The direct effects of changes in the exemption level or the cost of filing on the share of the population filing for bankruptcy are given by the following expressions:*

$$\partial p_i/\partial e = f_i(y^*) > 0,$$
$$\partial p_i/\partial c = -f_i(y^*) < 0.$$

The expressions above are of equal and opposite magnitude, implying that marginal changes in the cost of filing and the exemption level have similar effects, with the magnitude determined by the mass of marginal individuals who were indifferent to filing prior to the reform.

Next we can examine the relationship between the bankruptcy code (i.e., $e$ and $c$) and the cost of borrowing ($r_i$). As a benchmark, we assume that the credit market is perfectly competitive. We thus define the equilibrium interest rate $r_i$ implicitly by setting the amount recovered by lenders to be equal to the amount of borrowing: $R_i(r_i) = b_i$.[13] To calculate the effect of changes in $e$ and $c$ on interest rates, we first define $R(r_i)$:

$$\begin{aligned} R_i(r_i) &\equiv \int_0^{e+(1+r_i)b_i-c} (\max\{0, y-e\}) f_i(y)dy + \int_{e+(1+r_i)b_i-c}^{\infty} ((1+r_i)b_i) f_i(y)dy, \\ &= \underbrace{\int_e^{e+(1+r_i)b_i-c} (y-e) f_i(y)dy}_{\text{Recovered from bankruptcy filers}} + \underbrace{\int_{e+(1+r_i)b_i-c}^{\infty} b_i(1+r_i) f_i(y)dy}_{\text{Recovered from non-filers}}. \end{aligned}$$

[13] An important implicit assumption in this setup is that each credit score segment is priced separately and competitively, and individual types are fixed and do not respond endogenously to either market prices of the policy reform.

Using this expression, we can calculate the effect on the interest rate of a reform that changes either the exemption level ($de$) or the cost of filing ($dc$) by implicitly differentiating the equation $R_i(r_i) = b_i$. In the Online Appendix we derive the expressions for these two comparative statics ($dr_i/dc$ and $dr_i/de$), and we discuss the technical conditions for the two expressions to have equal and opposite sign, just as in Remark 1 above.[14] Next, using these two expressions we can derive an expression for the pass-through of the reform-induced change in bankruptcy filings into borrowing costs. In other words, we calculate $dr_i/dp_i$ in the case where the change in the probability of filing has been affected by a reform that changes either the exemption level or the cost of filing, which yields the following intuitive pass-through expression:

$$\frac{dr_i}{dp_i} = \frac{c/b_i}{1-p_i} = \frac{1+r_i}{1-p_i} \cdot \frac{c}{(1+r_i)b_i} > 0. \tag{1}$$

The full derivation of this expression is given in the Online Appendix. In words, this expression states that a bankruptcy reform that decreases bankruptcies (either by reducing exemptions or raising filing costs) will decrease the interest rate, and the magnitude of the decrease will be increasing in the change in the probability of filing. The ratio is scaled by $c/b_i$, where $c$ is the amount repaid to creditors by the marginal filer. This suggests that for bankruptcy reform to have a meaningful effect on interest rates, it must be the case that $c/b_i$ is not small. This requires that marginal filers repay a meaningful share of their debts. Intuitively, this expression shows that the pass-through of bankruptcy reform to interest rates requires that marginal filers must face substantial (financial, hassle, or stigma) costs of filing.[15]

In order to calibrate a realistic benchmark for the magnitude of pass-through, the last part of equation (1) separates the pass-through expression into a product of two terms. The first term is the ratio $\frac{1+r_i}{1-p_i}$, and the second term, $\frac{c}{(1+r_i)b_i}$, represents what creditors recover from the marginal filer, $c$, over the expected repayment of non-filers, $(1 + r_i)b_i$. In other words, this second term represents the counterfactual recovery rate for the borrowers who are deterred (at the margin)

[14]The Online Appendix shows that the comparative statics for both reforms are equal and opposite sign if we are willing to assume $F_i(e) \approx F_i(y_i^*)$. This is a reasonable assumption when the amount to be repaid $(1 + r_i)b \approx c$, that is, the amount to be repaid is close to the full cost of filing.

[15]There is suggestive evidence that the collective cost of filing for bankruptcy is high. White (1998), for instance, estimated that 15 percent of households could benefit from filing for bankruptcy at a time when just over one percent did. Similarly, Indarte (2018) finds evidence in support of relatively large "all-in" costs of filing for bankruptcy.

from filing for bankruptcy. We calibrate the first term using measures of $r_i$ and $p_i$ for each credit score segment using the Consumer Financial Protection Bureau Consumer Credit Panel (CCP) and Mintel Comperemedia (Mintel) data on credit card offers. To capture a more accurate measure of borrowing costs, we additionally scale up the credit card interest rates based on the Nelson (2018) estimates of the difference between interest charges and fee-inclusive borrowing costs for the most closely corresponding credit score bin.[16]

To calibrate the second term in equation (1), we calculate default rates for each credit-score segment using the CCP. Specifically, we calculate the probability that a newly-opened credit card loan goes into 60-day default or is charged-off within 12 months after opening. To use the average default rates in our calibration, we make two simplifying assumptions. First, we assume that the 60-day default rate can be used as a reasonable proxy for the recovery rate. This may be a reasonable assumption if the expected recovery rate is small conditional on a 60-day default, and the recovery rate is very high if individuals have not reached 60-day default. Second, we assume that the *average* default rate within a credit score segment corresponds to the default rate for the *marginal* individual deterred from bankruptcy by the reform. In our setting, this may be a reasonable assumption since we use relatively fine credit score segments in our main analysis – 10-point credit score bins, with credit scores running from 500 to 840. Table 1 presents the weighted average results of this exercise, split by prime and subprime.

As shown in the more granular Appendix Table A3, most of the pass-through estimates are fairly tightly bunched around the weighted-average pass-through estimate (though the very low credit score segments have somewhat lower pass-through estimates between 0.55 and 0.80). The limited amount of heterogeneity in the calibrated pass-through estimates is interesting, and reflects a kind of "balancing out" of declining default rates, bankruptcy probability, and interest rates across the credit score distribution. The limited heterogeneity is also reassuring for our empirical analysis, since our research design is based on comparing across credit score segments with different values of $dp_i$ and relating these changes in bankruptcy filing to changes in interest rates $dr_i$. As

[16]Specifically, we estimate the mark-up due to fees by calculating the ratio between the fee inclusive charges and interest charges from Table 6 (Pre-CARD Act Price Distribution on New Accounts) in Nelson (2018). We use the credit-score-bin mark-ups from accounts with 6-11 months of cumulative borrowing. The mark-ups are similar to those implied by the survey respondents in Stango and Zinman (2009).

Table 1: Benchmarking Interest Rate Pass-through

| (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|
| | | | | | | Pass-through Calibration | |
| Credit Score Segment | Population Share | Bankruptcy Rate | APR | APR w/ Fees | 60-Day Default Rate | APR | APR w/ Fees |
| Subprime | 26% | 3.1% | 13.4% | 23.1% | 31.5% | 0.80 | 0.87 |
| Prime | 74% | 0.3% | 10.2% | 12.8% | 1.1% | 1.09 | 1.12 |
| All | 100% | 1.0% | 11.0% | 15.5% | 9.1% | 1.02 | 1.05 |

*Notes:* This table reports calibration estimates of pass-through estimates split by prime and subprime consumers. See main text for details and Appendix Table A3 for analysis by 10-point credit score segment. Subprime includes all individuals with credit scores less than or equal to 620. The APR column comes from Mintel data on credit card offers, and the bankruptcy and 60-day default rates (in both cases measured over the next 12 months) comes from the Consumer Credit Panel (CCP). The pass-through estimate comes from combining the estimates in columns according to equation (1), using the default rate to proxy for one minus the recovery rate, which is the first term in equation (1).

a result, we interpret calibration estimates as providing a plausible magnitude that we can use to benchmark our empirical results against and also providing evidence against substantial treatment-effect heterogeneity across credit score segments. This means that the average effect that we estimate in our difference-in-difference regressions may be informative for a range of credit scores, not just those credit score segments most affected by the bankruptcy reform.[17]

We conclude by highlighting two simplifying assumptions of the model to keep in mind when interpreting the empirical results. First, like most models of consumer bankruptcy in the literature, we assume that individuals make rational bankruptcy filing decisions with full information. In addition to appearing through the cost of filing $c$, a relaxation of this assumption could allow borrowers to under-estimate the financial benefit of filing.[18] Second, the model assumes that the market for loans is perfectly competitive. With imperfect pass-through, the same effect of reform on filings will lead to smaller change in interest rates, since some of the incidence of the reform will reduce firm profits. As a result, pass-through results that are smaller in magnitude than our calibrated estimates could be interpreted as consistent with imperfect competition.

[17] However, given that even this simple model highlights potential for treatment effect heterogeneity, we also report results of two-way fixed effects models allowing for heterogeneous treatment effects in Appendix Table A8, using the two-way fixed effects estimator developed by de Chaisemartin and D'Haultfoeuille (2019).

[18] Finkelstein and Notowidigdo (2018) provide a framework which incorporates this type of misperception, which may be more useful for welfare analysis if behavioral biases are important.

## 4 Data

Our analysis relies on three main data sets: legal dockets for all consumer bankruptcies in 78 (of 94) United States bankruptcy courts; Mintel Comperemedia data on credit-card offers made to more than 2,000 consumers each month; and hospital-discharge records for over half a million individuals merged with a ten-year panel of their credit reports.

Data on consumer bankruptcy filings come from the Public Access to Court Electronic Records (PACER) system. This dataset includes more than three million filings from 78 bankruptcy courts during our sample period of 2004 through 2007, roughly 86 percent of all filings in the United States during that period. We cut off our sample at the end of 2007 throughout our analysis in order to mitigate the effects of the Great Recession on our results. We validate the data for each district by comparing the filings in the PACER records with the official totals published by the Administrative Office of the United States Courts (AOUSC).[19] Appendix Table A1 details the sample coverage by chapter and quarter-year, and Appendix Section B.1 provides more details on the sample.

To study pass-through to credit-market pricing, we use Mintel Comperemedia (Mintel) data on credit-card offers.[20] Mintel collects credit card offers from a representative sample of households in the United States, who are paid to send all direct-mail credit-card offers they receive to Mintel.[21] The data includes demographic information on the households (age of head of household, household composition), details on the credit-card offers (type of credit, interest rates, fees), and some limited credit measures (importantly, these include the same credit score observed in the CCP). Data is collected monthly and includes approximately 350,000 credit card offers (7,000 per month) and 100,000 individual-month observations (2,200 per month). Appendix Table A4 provides summary statistics on offers and Appendix Section B.2 provides more details on the sample.

To study the insurance value of the bankruptcy option, we analyze administrative hospital-discharge records from the California Office of Statewide Health Planning and Development for the

[19]The sample does not include the universe of bankruptcies because 13 districts did not grant fee waivers, and we drop 3 districts from the sample because the bankruptcies in the data do not match the AOUSC statistics.

[20]We do not observe bankruptcy filings in the Mintel data, so we additionally use the Consumer Financial Protection Bureau Consumer Credit Panel (CCP) to estimate the bankruptcy filing risk for each credit score segment, combining public-record snapshots with credit score archives to estimate prospective filing probabilities.

[21]The sample is representative of United States credit card holders. According to Fulford and Schuh (2015), roughly 75 percent of American households hold at least one credit card.

universe of uninsured hospitalizations (and approximately 20 percent of individuals hospitalized with insurance) between 2003 and 2007. Our sample links hospitalized individuals to their panel of credit reports spanning the years 2002 to 2011. To isolate unexpected hospitalizations, we restrict the sample to individuals ages 25 to 64 who are hospitalized for non-pregnancy-related reasons and have not previously been to the hospital in the last three years. Appendix Table A9 provides pre-hospitalization summary statistics and Appendix Section B.3 provides more details on the sample.

## 5 Effect of BAPCPA on Filings in the Short and Medium Run

### 5.1 Effect on Total Filings

Figure 1 plots the total number of consumer bankruptcy filings in the PACER sample by week from January 2004 through December 2007. The most striking feature of Figure 1 is the dramatic rush to file after BAPCPA was signed but before the bankruptcy code was changed. In the five weeks before the law was implemented, from September 12$^{\text{th}}$ through October 16$^{\text{th}}$, the filing rate increased dramatically. In the final week before the implementation of the law, more than 400,000 households declared bankruptcy, roughly 13 times the typical weekly caseload.

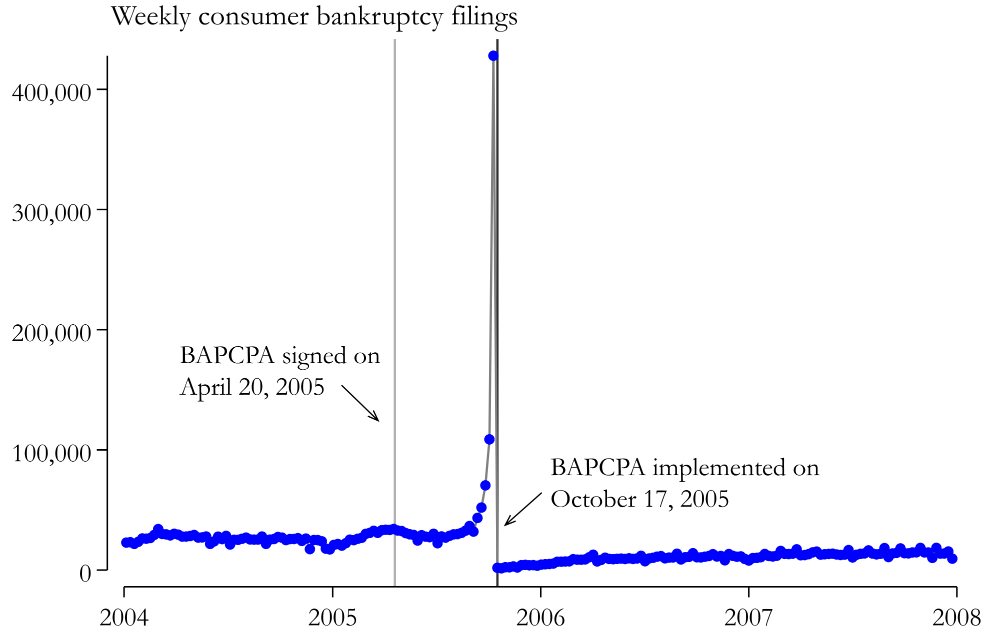
To quantify the number of excess filings before implementation and to test whether, on net, the law led to a reduction in bankruptcies, we adapt "excess mass" methods from the tax-notch literature (e.g., Chetty et al., 2011; Kleven, 2016) to generate a counterfactual time-series in the absence of the changes to the bankruptcy code. We fit the following regression to the weekly filing count in the period before BAPCPA was passed by the Senate (March 10, 2005):

$$\text{Filings}_t = \gamma t + \tau_m + \varepsilon_t. \tag{2}$$

Here, $\text{Filings}_t$ are nationwide filings in week $t$, $t$ is a linear time trend, and $\tau_m$ are calendar-month fixed effects. Appendix Table A2 presents results when we additionally control for the national unemployment rate and show robustness to alternatively fitting the counterfactual time-series through the passage of the bill in the House (April 14, 2005) and its signing into law (April 20, 2005). The results are qualitatively similar across the specifications.

We use Equation 2 to predict the counterfactual number of filings each week for the full sample

Figure 1: Time-Series of Bankruptcy Filings



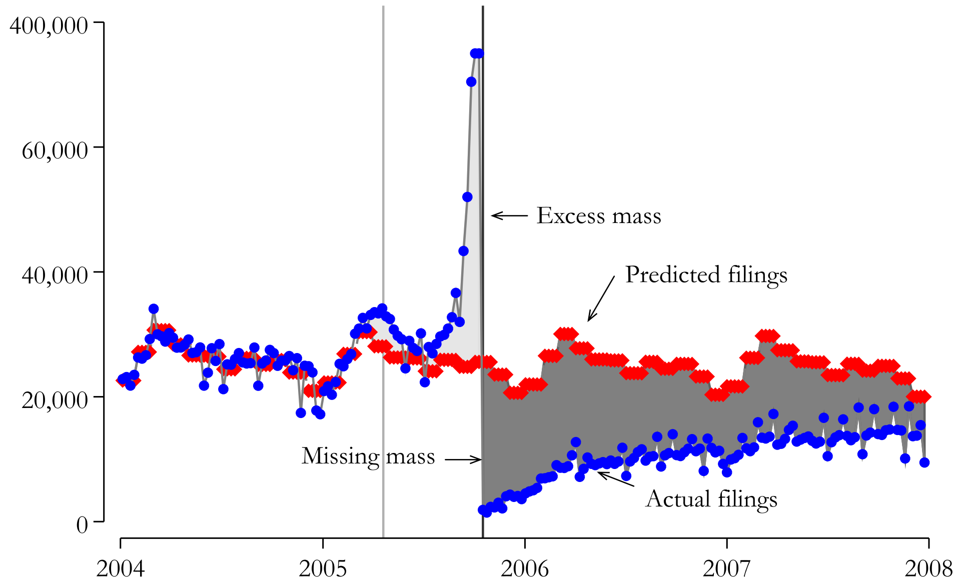

*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007. Each dot in the figure represents the total count of filings for that week.

period and calculate the sum of the difference between the predicted and actual filings for each week. Figure 2 presents this exercise by plotting the time-series of bankruptcy filings against the estimated counterfactual. As expected, the predicted filings closely match actual filings before the passage of the law. Actual filings diverge from the predicted time-series in September of 2005 in advance of the pre-BAPCPA filing deadline in mid-October. An excess of more than 750,000 households filed for bankruptcy between March 10, 2005 and October 17, 2005 relative to the counterfactual time-series. To calculate the net effect of BAPCPA on filings, we account for the filings which were intertemporally substituted before the implementation of the law.

Table 2 presents the difference between the predicted and the realized number of filings from the bill's passage in the Senate through the end of 2007. Column 1 presents the predicted number of filings for the 30 week period ending in the index date. Column 2 presents the difference between the realized and predicted filings for the same period. Column 3 presents the cumulative net difference in filings from the counterfactual time-series.

The more than 750,000 excess filings ahead of implementation suggest that debtors and their attorneys anticipated the changes to the bankruptcy code to be significant. Due to the mandated

Figure 2: Excess and Missing Mass of Bankruptcy Filings




*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007. The total count of filings for each week is plotted against the predicted number of filings for the week. The predicted number of filings are the result of estimating equation 2 on the total count of filings from January 2004 through the day that BAPCPA was passed by the Senate (March 10, 2005). The two data points before implementation of BAPCPA are censored in this figure: there were 108,745 filings during the week that began on October 3, 2005 and 427,947 filings during the week that began on October 10, 2005.

waiting period betweeen filings, whose who file for bankruptcy must do so because the benefits of filing today must exceed the loss of the option to file at another point in the future.[22] For debtors who rushed to file before the new bankruptcy code went into effect, we can infer that the benefit from filing for bankruptcy under the previous system exceeded the option value of bankruptcy under the new system.

On net, the decline in filings after implementation exceeds the pre-implementation increase in filings by July of 2006. The temporary "rush to file" effect was quickly overwhelmed by the sustained reduction in filings under the new bankruptcy regime. As of the end of 2007, 114 weeks after the implementation of BAPCPA, 1,077,679 filings were deterred.[23] Appendix Figure A1

[22]Before BAPCPA, filers needed to wait 6 years after a Chapter 7 bankruptcy before filing again. BAPCPA increased that waiting period to 8 years, a change that Appendix Figure A4 demonstrates was binding for the small share of filers who file more than once.

[23]To calculate a confidence interval for the estimated net change in filings, we implement the bootstrapping procedure described by Chetty et al. (2011). This leads to a 95-percent confidence interval with an upper bound of 1,125,242 and a lower bound of 1,034,709. Since we observe the full sample of bankruptcy filings, the standard errors reflect error due to misspecification of equation (2) rather than sampling error.

Table 2: Difference between Realized and Predicted Filings

| | | (1) | (2) | (3) |
|---|---|---|---|---|
| Weeks relative to implementation | Index Date | Predicted Filings | Realized Difference | Cumulative Net Difference |
| -30 | March 21, 2005 | | | |
| 0 | October 17, 2005 | 911,656 | 762,192 | 762,192 |
| 30 | May 15, 2006 | 879,729 | -656,283 | 105,909 |
| 60 | December 11, 2006 | 857,796 | -481,442 | -375,533 |
| 90 | July 9, 2007 | 889,823 | -445,607 | -821,140 |
| 114 | December 24, 2007 | 659,619 | -256,539 | -1,077,679 |

*Notes:* This table presents a running sum of the net change in filings due to BAPCPA: the difference between actual bankruptcies observed each week and the number of bankruptcies that would have been predicted based on the counterfactual by estimating equation 2 from the beginning of the sample until BAPCPA was approved by the Senate in March of 2005. Index date for each row refers to the end of the 30 weeks period presented. The overall numbers are inflated to reflect the nation as a whole, based on our PACER sample coverage (see Appendix Table A1).

presents a similar exercise, but extending the pre-period back to 2002. This generates an estimate of -1,109,094 which is consistent with the estimates using the 2004 through 2007 period.

### 5.2 Effect on Choice of Chapter

Bankruptcy reform clearly decreased the overall number of filings, but the introduction of the means test also sought to shift more filings from "fresh start" bankruptcies (Chapter 7) to repayment-plan bankruptcies (Chapter 13). Appendix Figure A2 plots the time-series for total filings separately for Chapter 7 and Chapter 13. The time-series patterns are not markedly different across the two chapters, as we might expect if the primary impact of the reform was a means-test-driven shift in the chapter of filings; however, the decline in filings is larger among Chapter 7 filings. The share of filings that were Chapter 13 remained persistently higher after the reform, as is clear from Appendix Figure A3. Appendix Table A2 estimates the net change in overall filings through 2007 by chapter. Through 2007, both Chapter 7 and Chapter 13 filings declined substantially. This result could be consistent with the law's intention to shift high-income filers to Chapter 13, which Section 7 examines directly.

## 6 Effects of Bankruptcy Filing Risk on Interest Rates

### 6.1 Pass-through to Interest Rates

In determining interest rates, creditors must predict the expected repayment rates on the credit offered. A key input for determining repayment rates is the probability an individual will discharge their debt through bankruptcy. Filing for bankruptcy either reduces repayment to zero under Chapter 7 or restructures the amount to be repaid under Chapter 13. As suggested by the model in Section 3, reducing the generosity of the bankruptcy system ought to increase prospective repayment rates, and thus decrease the cost of lending. We expect this decrease in the cost of lending to be passed-through to borrowers in the form of lower interest rates. We test this hypothesis by estimating the pass-through to interest rates on credit-card offers.

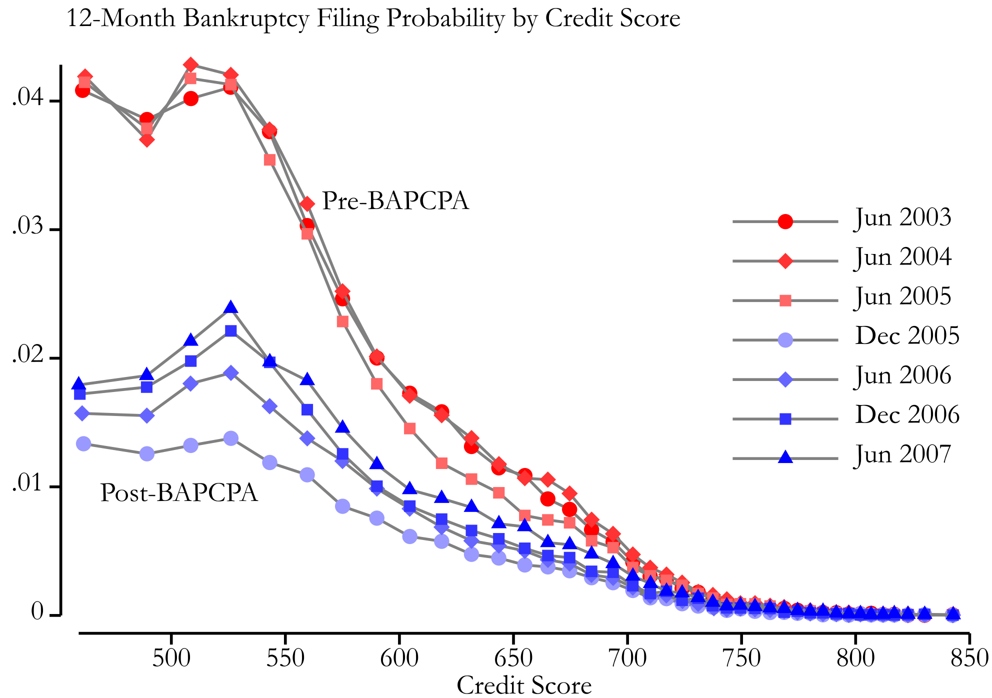
We focus on credit-card offers for three reasons. First, credit cards are the most common method of borrowing in the United States. Roughly 75 percent of Americans have at least one credit card (65 percent of whom carry a balance) and total revolving debt was over $800 billion for most of our sample period (Fulford and Schuh, 2015). Second, because credit-card debt is not collateralized, it is the most likely to be discharged in bankruptcy and thus the most responsive type of credit to changes in the bankruptcy code. Third, the Mintel dataset provides a clean measure of the cost of credit supplied to households in the credit card market, allowing us to overcome measurement challenges associated with other data sources.[24]

We are interested in identifying the change in borrowing costs ($dr$) for a given change in bankruptcy filing risk ($dp$). Our empirical approach to estimate pass-through is motivated by the observation that the bankruptcy risk of a potential borrower varies substantially by credit score. This is evident in Figure 3, which plots the probability that borrowers across 40 equal-sized credit score segments file for bankruptcy over the next 12 months for each available CCP observation. Filing risk conditional on credit score bin decreased significantly after the new bankruptcy code was implemented.[25]

[24] For instance, credit bureau data do not include prices. Other datasets, such as the National Mortgage Database, offer information on prices conditional on loan take-up.

[25] This was not driven by changes in the credit scoring formula across the sample period. We use the same credit score throughout the period and there were no major changes to the standard commercially available credit scoring formulas over this period.

Figure 3: Probability of Bankruptcy




*Notes:* The sample are individuals in the CFPB CCP. This figure plots a binscatter of the share of individuals who file for bankruptcy within the next 12 months. Each point represents the 12-month prospective filing rate for one of 40 equal-sized credit score segments from the time of the credit report observation.

To parameterize the change in the probability of filing for bankruptcy, we define $\delta_b$ as the difference between the post-BAPCPA filing probability and the pre-BAPCPA filing probability for each 10-point credit score segment. We estimate the change in the *prospective* bankruptcy filing risk, by comparing the 12-month bankruptcy filing rates for each credit score segment before and after the reform. In particular, we take the average 12-month filing rate for each available quarter of the CCP before and after the reform was implemented, and define $\delta_b$ as the change in filing risk.[26] Appendix Table A5 presents the estimates of $\delta_b$.[27] We use the change in bankruptcy filing probability, $\delta_b$, as our parameterization of the change in bankruptcy filing probability we describe

[26]We use the prospective 12-month filing rate for each of the time periods plotted in Figure 3 and take the difference in the average between the pre-period observations (observed quarterly from September 2003 to September 2005, missing December 2003) and the post-period observations (observed quarterly from December 2005 to December 2007).

[27]BAPCPA also implemented a longer waiting period between filings, increasing the minimum time between Chapter 7 filings from six to eight years. Appendix Figure A4 demonstrates that this was binding for a small share of repeat filers. Given the large bunching documented in Section 5, it is useful to decompose any change in the probability of filing for bankruptcy that is attributable to a change among those who are eligible and an increased "incapacitation" effect among those no longer eligible to file. Appendix Figure A5 plots the change in filing probability, $\delta_b$, separately by whether an individual has a prior filing and shows that the decline in filing is driven almost entirely by a change in the probability of filing amongst those who are eligible.

in section 3 (i.e., $\frac{dp}{dc}$). This allows us to estimate the response of interest rates with respect to bankruptcy filing rates, which maps to the comparative static of interest (i.e., $\frac{dr}{dp}$) we derived.[28]

Our empirical strategy amounts to a comparison of different credit-score segments. We compare the changes in the interest rates offered to segments that experienced large declines in bankruptcy filing risk to the interest rates offered to segments that experienced little change in bankruptcy filing risk. That comparison isolates the object of interest ($\frac{dr}{dp}$) so long as the change in bankruptcy risk across credit-score segments is not correlated with other time-varying determinants of interest rates. That assumption would be violated if, for instance, the credit-scoring formula itself were changing during this time period. Reassuringly, however, the credit-scoring formula was unchanged during this time period and the distribution of credit scores was quite stable (Appendix Figure A6).

That identification assumption also requires parallel trends. That is, the assumption requires that, absent BAPCPA and conditional on these controls, the pre-period differences in offered interest rates across segments would have evolved along parallel trends. That assumption would be violated if, for example, credit-score segments that experienced larger declines in bankruptcy filing probability also experienced larger changes in interest rate offers for reasons beyond the factors we can include as controls. For instance, lenders were increasingly focused on the subprime mortgage market during this time period, and so that portion of the market may have been evolving along a very different path. We address this concern by showing that the estimates below are robust to the inclusion of subprime- and prime-specific time trends.

The Mintel data is a repeated cross-section and the level of observation is a credit-card offer. In our main specifications, we include lender-specific fixed effects to absorb differences across lenders, credit-score-segment fixed effects to absorb time-invariant differences in prices across credit-score segments, and fixed effects for card category and application type (e.g., pre-approved or invitation to apply). The estimating equation for the event study is:

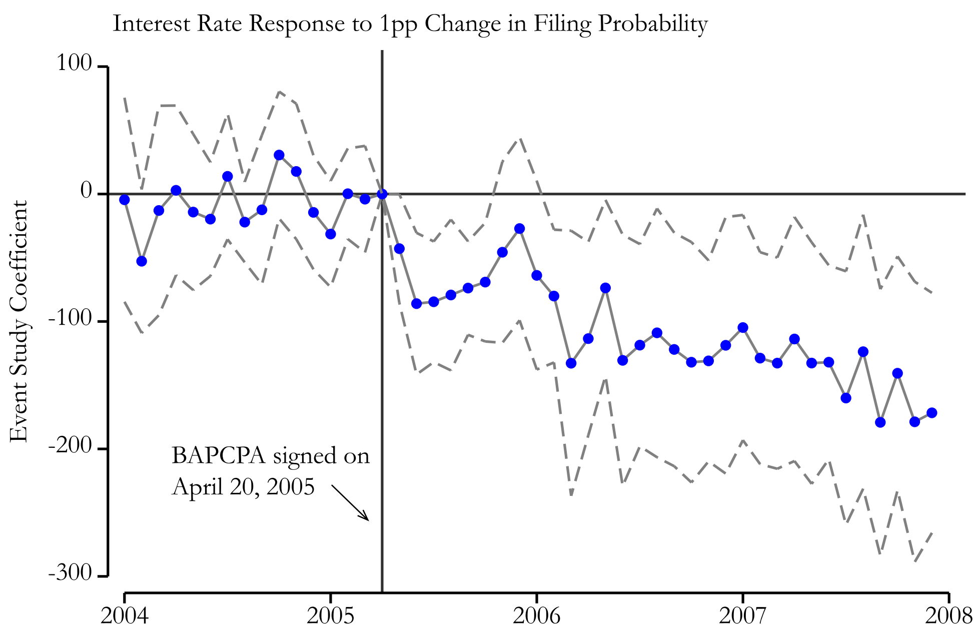
$$y_{ijt} = \beta_0 \delta_b + \sum_{2004m1}^{t=2007m12} \lambda_t (\delta_b \times \tau_t) + \nu_j + \chi_i + \phi_b + \varepsilon_{ijt}. \tag{3}$$

[28]One can also regress equation 4 using only the pre-reform bankruptcy filing risk. Appendix Table A6 presents additional regressions that use only the pre-BAPCPA risk of filing for bankruptcy in lieu of the difference. As expected, the results are consistent with estimates using the change in filing probability.

The dependent variable $y_{ijt}$ is the interest rate of offer $i$ for credit-score segment $j$ in month-year $t$. The variable $\delta_b$ is the "treatment" (the difference in the propensity to file before and after passage of BAPCPA) and $\tau_t$, $\nu_j$, $\chi_i$, and $\phi_b$ are fixed effects for each month and year combination, lender, credit score segment, and other offer features respectively.[29] We two-way cluster standard errors by credit-score segment and lender (Cameron et al., 2011).

Our primary outcome is the regular annual percentage rate (APR). Appendix Figure A7 presents the time-series of the regular interest rate, split by prime and subprime and adjustment for the prime rate. We weight all credit-card offers using Mintel-provided weights designed to make the sample representative of the overall mail volume of each campaign. This allows us to estimate effects that are representative of the credit card market as a whole.[30] In the appendix, we also examine the adjusted APR, which follows Gross et al. (2016) in adjusting for whether the credit-card offer has an introductory "teaser" rate by taking a weighted average of the introductory interest rate and the regular interest rate over the first 12 months after origination.

Figure 4: Effect of Decline in Filing Probability on Offered Interest Rates




*Notes:* The sample is credit card offers made between January 2004 and December 2007 included in the Mintel data. The points represent estimates of the $\lambda_t$'s in equation 3. The dashed lines provide 95% confidence intervals for each point. The dependent variable is the rate spread for regular offered interest rate.

[29]The other features of the credit-card offers are: card category, application type, and state of residence.

[30]Mail volume represents the effective weight on each mail piece. For instance, the sum of all mail volume weights in the full Mintel data for *Chase Freedom* cards in a given month should equal the total mailings for *Chase Freedom* in that month, nationwide. In practice, the weights do not meaningfully affect the coefficients.

Figure 4 plots the coefficients of interest, $\lambda_t$ for each month $t$, when the regular APR spread is the dependent variable in equation (3). By allowing the $\lambda_t$'s to evolve flexibly over time, this regression allows us to assess the assumption that interest rate offers were evolving along parallel trends before the passage of BAPCPA. Further, by refraining from imposing any ex-ante restrictions on when interest rates should change, we can use the time pattern to gauge whether any changes in interest rates appear after the passage of BAPCPA but before the law's implementation.

In equilibrium, we would expect any changes to the bankruptcy code and interest rates to also affect borrowing behavior. We expect offered interest rates to respond immediately to anticipated changes in bankruptcy filing probability, while any secondary effect on borrowing would take longer. Therefore, we are particularly interested in whether there exists a break in the evolution of interest rate offers when BAPCPA was signed into law. We focus on the timing of passage (rather than implementation) here because the bankruptcy code considers debts incurred in the months just before filing as non-dischargeable; therefore, creditors could safely assume new lines of credit opened between passage and implementation would not be discharged before the new bankruptcy code took effect.

Figure 4 suggests a decline in offered interest rates ahead of BAPCPA's implementation. While interest rates evolved similarly for the credit-score segments affected and unaffected by the reform throughout the pre-period, we observe a sharp drop in the $\lambda_t$'s following the passage of BAPCPA in April 2005. The decline in interest rates among the credit-score segments who experienced a decline in the probability of filing for bankruptcy is stark and persistent—the interest rate spread drops immediately upon passage of BAPCPA and remains below the pre-period level through the post-period.

Motivated by the pattern in Figure 4, we estimate a difference-in-difference regression to quantify $\frac{dr}{dp}$. Specifically, we estimate the change in offered interest rates for a one-percentage-point decline in the 12-month probability of filing for bankruptcy:

$$y_{ijt} = \beta_0 \delta_b + \beta_1 \delta_b \times \text{post} + \tau_t + \nu_j + \chi_i + \phi_b + \varepsilon_{ijt}, \tag{4}$$

where, again, $\delta_b$ is the difference in the probability of filing before and after the passage of BAPCPA and $\nu_j$, $\chi_i$, and $\phi_b$ are indicators for lender, offer features, and credit-score segment. In lieu of the month-year indicators interacted with $\delta_b$, we simply interact $\delta_b$ with a "post" indicator for the offer coming after the Senate passage of BAPCPA. Offers are weighted by mail volume and standard errors are two-way clustered by credit-score segment and lender.

Table 3: Pass-through of Change in Bankruptcy Filing Probability to Interest Rates

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Dependent variable: | | | | | | | | |
| Post-BAPCPA $\times \delta_b$ | -94.6* | -78.7 | -101.3*** | -60.4** | -59.1** | -66.7** | -66.6** | -66.4** |
| | (50.7) | (48.4) | (35.7) | (27.2) | (24.6) | (29.3) | (29.0) | (28.9) |
| $R^2$ | 0.15 | 0.19 | 0.43 | 0.43 | 0.52 | 0.55 | 0.55 | 0.55 |
| Month-by-Year FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Score Bin | | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Lender FEs | | | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Subprime $\times t$ | | | | ✓ | ✓ | ✓ | ✓ | ✓ |
| Card Category | | | | | ✓ | ✓ | ✓ | ✓ |
| Application Type | | | | | | ✓ | ✓ | ✓ |
| State FE | | | | | | | ✓ | ✓ |
| Census Division $\times$ Year FE | | | | | | | | ✓ |
| $N$ | 391,279 | 391,153 | 391,153 | 391,153 | 390,975 | 390,975 | 390,975 | 390,381 |

*Notes:* The sample consists of credit card offers made to households from January 2004 through December 2007. All columns report effects based on OLS estimates of equation 4. The outcome variables are the interest rates on credit card offers, adjusted for the prime rate. Standard errors (two-way clustered by credit score segment and lender) are in parentheses. Offers are weighted by the mail volume of the campaign. Asterisks indicate significance at the 1 percent (***), 5 percent (**), and 10 percent (*) level, respectively.

Table 3 presents the coefficients of interest, $\beta_1$, from Equation 4. The baseline specification in column 3 includes month-by-year, credit-score segment, and lender fixed effects.

In order to control for other factors which may have changed differentially by credit-score segment and correlated with $\delta_b$, such as the expansion of the subprime market, we include prime- and subprime-specific time trends starting in column 4. The main coefficient declines by roughly forty percent with the inclusion of time trends. That change in the coefficient is unsurprising in that the subprime market was expanding dramatically during this time period. After accounting for prime- and subprime-specific time trends, however, the estimates are very stable through the incremental inclusion of card category, application type, state-specific fixed effects, and Census-Division-by-year fixed effects.

Our preferred estimate of the relationship between interest rates and bankruptcy filing risk comes from the most-demanding specification in Table 3: a 67-basis-point decline in the interest rate for each 1-percentage-point decline in the 12-month prospective bankruptcy filing risk. Borrowers who were in a credit-score segment whose bankruptcy risk fell by two percentage points (a relatively common decline observed in Figure 3) received interest rate savings of approximately 134 basis points. Based on Mintel data on interest rate offers during the pre-period, this difference in mean interest rate offers is roughly equivalent to moving from a 600 to 650 credit score.[31]

We can use the pass-through expression derived in Section 3 to assess the magnitude of these estimates. The perfect-competition benchmark for $\frac{dr/dc}{dp/dc}$, above, is 105 basis points for each one-percentage-point change in filing risk. The estimates of Equation (4) are of similar magnitude to this benchmark. The estimate in column 4 of Table 3 represents pass-through of approximately 64% of the perfect-competition benchmark. While this exercise is not a conclusive statement on the competitiveness of credit card markets, it does suggest that our estimates of pass-through are of reasonable magnitude and consistent with some competitive pricing pressure as well as imperfect competition.

These decreases in offered interest rates benefit consumers to the degree that they take advantage of the lower rates. Consumers may not be especially strategic in doing so. Stango and Zinman (2013) and Woodward and Hall (2012) document substantial dispersion in borrowing costs that appear to be driven by consumers partaking in too little price-shopping. Similarly, Gathergood et al. (2017) document that consumers tend to repay their credit-card debt based on heuristics rather than by minimizing borrowing costs. Nevertheless, as long as consumers borrow using unsecured credit, the decrease in interest rates as a result of BAPCPA are likely to have provided them with considerable savings.

[31] Appendix Table A7 presents a similar analysis for adjusted interest rates and also suggests a large decrease in offered prices. The estimated effect for adjusted interest rates, however, is smaller than the effect for the regular interest rate. Ru and Schoar (2016) show that less financially sophisticated households are more likely to be offered teaser rates. Nevertheless, this contrast suggests that the majority of the change in prices was driven by decreases in the regular interest rate rather than an expansion of teaser rates.

### 6.2 Heterogeneity Analysis

The pass-through calculation in this section is based on the average effect reported in Table 3. Given the potential for treatment effect heterogeneity highlighted by the model, we re-estimated our main specification using a recently developed two-way fixed effects estimator that allows for heterogeneous treatment effects (de Chaisemartin and D'Haultfoeuille, 2019). Reassuringly, we find similar results using this alternative estimator, and we also find no "negative weights" going into the average treatment effect. This is consistent with the calibrations in Section 2, which show fairly tight bunching around the weighted-average pass-through estimate across credit score segments.

We perform one additional heterogeneity analysis by examining how the response to the reform varies by lender. Figure 4 and Table 3 demonstrate that lenders lower the interest rates on new credit card offers in anticipation of a decline in the generosity of bankruptcy. Examining heterogeneity in the responses of lenders can serve to both clarify the mechanism behind those lower interest rates and reveal some information about the structure of the subprime credit card market during our sample period. The Mintel data include the name of the lender that issued each offer, which we partial out in regression equation (4). To understand how the responses to the change in bankruptcy law varied by lender, we run equation (4) separately for each lender.

Appendix Figure A9 plots the lender-specific estimates on the vertical axis against the share of the firm's offers that are made to subprime consumers on the horizontal axis. The size of each circular marker corresponds to the number of credit card offers extended to subprime consumers over our sample period. The effects we estimate in Table 3 are largely driven by the three prominent subprime lenders: Capital One, HSBC, and Washington Mutual. Capital One and HSBC, in particular, are well-known to be leaders in the expansion of credit to subprime consumers while HSBC was additionally on the frontier of expanding credit to those with recent bankruptcy filings, which suggests a familiarity with the relationship between the bankruptcy system and repayment behavior (Jurgens and Wu, 2007).

## 7 Effects of BAPCPA on Targeting and the Insurance Value of Bankruptcy

The previous section documents how a decrease in bankruptcy filings passed-through to lower interest rates. This benefit to borrowers came at the cost of worsening the value of the bankruptcy option. How we weigh these impacts depends on *which* potential bankruptcy filers were deterred. Bankruptcy reform had the explicit goal of deterring "abusive" filings from higher-income filers who could otherwise repay their debts. In this section, we evaluate whether the means test was successful at shifting the distribution of filers away from lower-insurance-value bankruptcy filings, and estimate the effect of the reform on the likelihood that a negative financial shock is insured by bankruptcy.

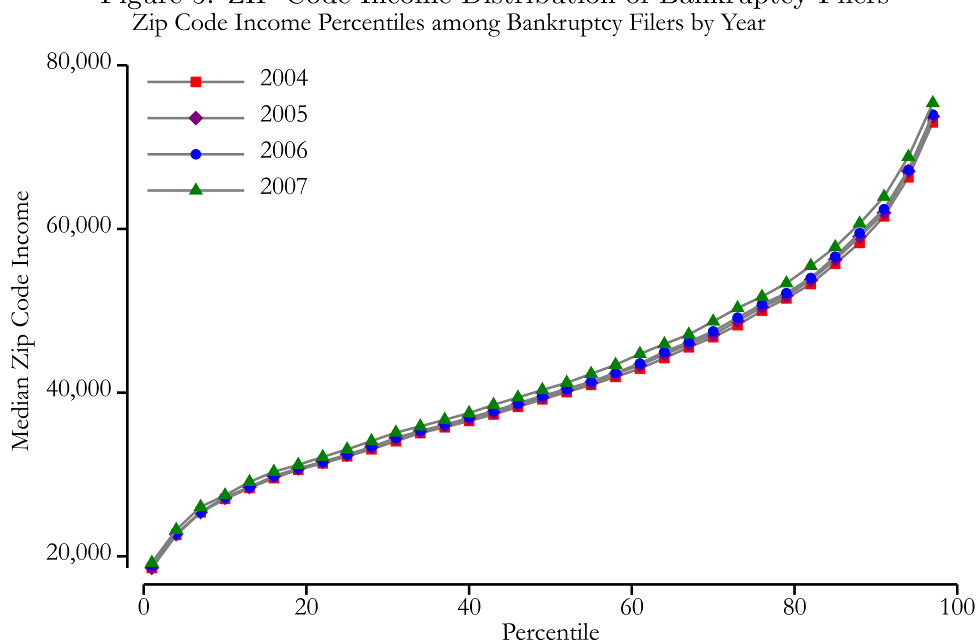
### 7.1 Characteristics of Filers

A key goal of bankruptcy reform was to deter high income filers from accessing bankruptcy relief "opportunistically;" lawmakers referenced the income-based means test as the "heart of the bill" House Report (2005). By excluding households with income above the state median from the option to liquidate their debts, the law intended to target the bankruptcy code's most-generous provisions to lower-income filers. If the means test was an important force altering the composition of filers, we would expect to see the income of the average filer decrease.

While much attention has been given to the means test, the reform's collective impact on the composition of filers is ambiguous *a priori*. BAPCPA also made a number of additional changes to the bankruptcy filing process which collectively increased both the hassle costs of filing (through mandated credit counseling and financial management courses) and the liquidity requirements to file (through increased filing and attorney fees). The increased liquidity requirements to file, especially, may have deterred lower-income filers. The effect of hassle costs on the composition of filers is ambiguous, and depends on both filers' opportunity cost of time and their ability to navigate the requirements or to pay an attorney to help them do so.

The overall impact of the reform on the income of filers depends on the relative deterrent effects of these provisions across the income distribution. A cursory reading of the decline in filings, and the decline in Chapter 7 filings in particular, suggests the means test may have been effective in achieving its stated goals of deterring higher-income filers. This interpretation is belied by a closer examination of the composition of filers. We examine how the income of filers evolved through the reform by merging the filer's ZIP Code in the PACER sample with the median income for that ZIP Code from the 2000 Decennial Census.

Figure 5: ZIP Code Income Distribution of Bankruptcy Filers




*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007, matched with the ZIP Code median household income measured in the 2000 decennial census. The figure plots the percentiles of ZIP Code median household income among filers for each year of 2004 through 2007.

Figure 5 plots the full distribution of median ZIP Code income among filers for each year from 2004 through 2007. The distributions are strikingly similar—percentiles are virtually on top of each other through the 60$^{\text{th}}$ percentile, at which point the post-BAPCPA distribution of filers drifts slightly upward. As is clear from the figure, there is no stark change in the composition of filers.[32] This suggests that the reform did not change the composition of filers by very much in terms of income (at least with respect to their ZIP Code), and, if anything, average income appears to creep upward after the reform. This surprising null effect ought to be explored by future research that examines how BAPCPA changed households' bankruptcy decision rules.[33]

[32] Of course, income also varies within ZIP Codes and it is possible that there were large within-ZIP-Code changes in the incomes of filers. Nevertheless, we view it as unlikely that the ZIP Code income measure is masking large means-test-driven shifts in the income distribution of filers. This is consistent with anecdotal reports from bankruptcy attorneys (Littwin, 2016) and other evaluations of the reform (Ashcraft et al., 2007; Albanesi and Nosal, 2018).

[33] Several papers have used the means test as variation with which to study the effects of BAPCPA and bankruptcy

Inspecting differences in the composition of filers before and after the reform can reveal how the self-targeting properties of the bankruptcy system changed. However, to measure changes in the insurance value of bankruptcy, we need to test how individuals facing the same negative financial shocks were able to access bankruptcy before and after BAPCPA. The next section takes such an approach.

### 7.2 Effect of BAPCPA on the Insurance Value of Bankruptcy

The results above demonstrate that BAPCPA deterred filings. To evaluate the cost of these deterred filings, we next ask how expense shocks (specifically, health shocks requiring hospitalization) were insured by bankruptcy before and after the reform. In the debate over bankruptcy reform, many expressed concern over how reform might affect the insurance value of bankruptcy. Warren (2005) argued against BAPCPA because the means test would "treat all families alike. . . A person who had a heart attack is treated the same as someone who had a spending spree at the mall."

This distinction between bankruptcies driven by medical costs and bankruptcies driven by discretionary consumption is present in life-cycle models of the bankruptcy decision.[34] Livshits et al. (2007) demonstrate that the existence of expense shocks, such as medical costs, can make "fresh start" (Chapter 7) bankruptcy regimes welfare-increasing despite increasing the cost of borrowing. Particularly when markets are incomplete, bankruptcy may be the only mechanism by which an individual can insure some negative events. We thus seek to estimate whether *specific expense shocks* were insured by bankruptcy, before and after the reform. We test the likelihood that individuals experiencing hospitalization shocks, before and after changes to the bankruptcy code, declare bankruptcy to obtain debt relief.

We study the universe of uninsured hospitalizations between 2003 and 2007 in California, where approximately 20 percent of residents lacked insurance during that time (California Healthcare Foundation, 2010). For comparison, we study insured hospitalizations in parallel. The insured include adults ages 25–64 hospitalized with either private or Medicaid coverage. Each hospitaliza-

more generally. See, for example, the work of Chatterjee et al. (2007), Li et al. (2011), Mahoney (2015), Mitman (2016), and Parra (2018)

[34] Divorce, job loss, and unplanned pregnancies are additional shocks that are discussed as relevant to the welfare implications of the bankruptcy code (Livshits et al., 2007; Fay et al., 2002; Keys, 2018), though medical expenses are often pointed to as the most "blameless."

tion is linked to the patient's credit reports, which we observe at the start of each year from 2002 through 2011. We limit the sample to those who were not hospitalized in the three years prior to their hospitalization to isolate "health shocks."[35]

In the ideal experiment, we would randomly assign different bankruptcy regimes to otherwise-identical individuals experiencing a health shock. One might be concerned that the composition of uninsured hospitalizations might be different before and after the reform.[36,37] By isolating health shocks for individuals who have not been hospitalized in the previous three years, we come close to approximating this experiment. Hospitalizations are much less likely to be anticipated than other sources of health insurance demand like chronic conditions requiring outpatient care. To further address this concern, we re-weight the two sets of hospitalizations on their observable characteristics, though it has little effect on our estimates. We use propensity score matching to reweight those hospitalized in each period in order to match them on age, sex, race, zip code household income, whether the hospitalization was for a chronic condition, and on the major diagnostic category. Appendix Table A9 presents summary statistics by insurance status and hospitalization period.

Following Dobkin et al. (2018a), we estimate event-study regressions, additionally splitting the sample by whether the hospitalization occurred under the pre- or post-BAPCPA bankruptcy regime. We define the pre-BAPCPA period to be January 2003 through December 2004 and the post-BAPCPA period to be from October 2005 through December 2007.[38]

We define event time $m$ as the number of months relative to the hospitalization, which occurs at $m = 0$. Omitting the month prior to the hospitalization ($m = -1$) and including calendar-year-specific fixed effects, we specify a non-parametric event-study regression to estimate the evolution

[35] The precipitating event will be a hospitalization, but estimated effects of the hospitalization will include all sequelae (subsequent hospitalizations, poor health, lost earnings).

[36] As observed by Mahoney (2015), changes in the bankruptcy code also changes the incentives for individuals to purchase health insurance. This generates a concern that uninsured health shocks would be differentially selected before and after BAPCPA. Two patterns in the data ameliorate these concerns. First, the means test does not appear to be the primary driver of the change in bankruptcy filings. Second, at least in a coarse examination of the data, the share of Californians without health insurance was broadly unchanged over our sample period (California Healthcare Foundation, 2010). The estimated share of individuals lacking health insurance in California in each year from 2003 through 2007 was 19.1%, 19.5%, 19.9%, 19.6%, and 19.4%.

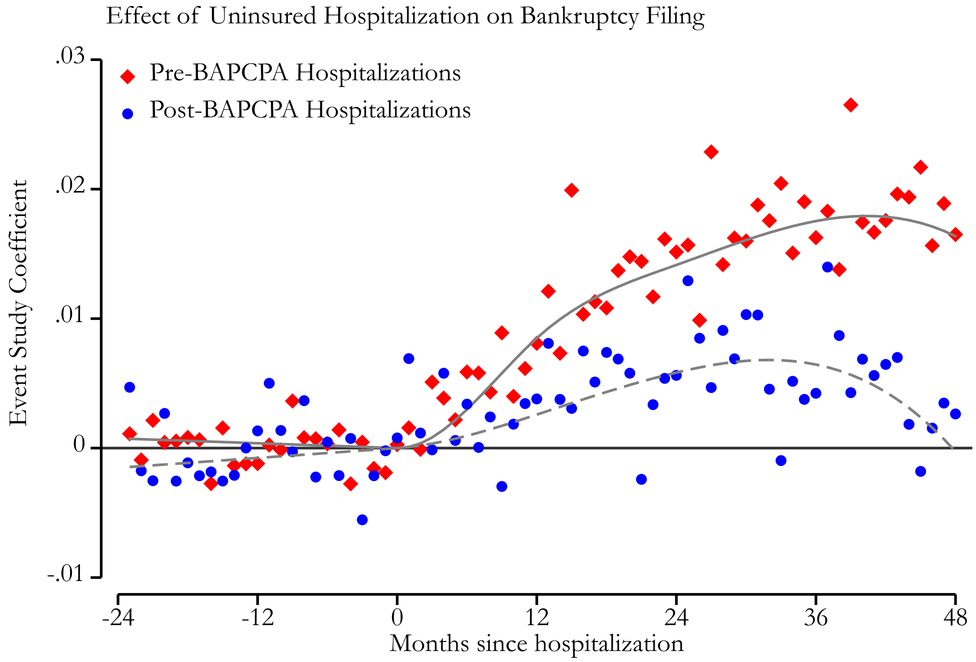
[37] A related concern is that the financial consequences of a hospitalization may vary over time. California implemented the "Hospital Fair Pricing Act" in 2007 which required hospitals to offer discounts or charity care to individuals making less than 350% of the Federal Poverty Line. Appendix Figure A13 plots the implied effects by year of hospitalization, and shows the results within the post-period do not look dramatically different between 2006 and 2007 hospitalizations.

[38] We exclude hospitalizations occurring between January 2005 and September 2005 to avoid those most likely to coincide with the rush-to-file in October 2005.

of the outcome variable preceding and following the hospitalization:

$$y_{it} = \gamma_t + \mathbb{1}\{\text{Pre-BAPCPA}\}\left(\sum_{m=-24}^{-2}\mu_m + \sum_{m=0}^{48}\mu_m\right) + \varepsilon_{it}. \quad (5)$$

Figure 6: Effect of Hospitalization on Bankruptcy Filing




*Notes:* The sample is individuals ages 25-64 who are hospitalized without insurance in California, additionally split by the timing of the hospitalization (January 2003 through December 2004 for the pre-BAPCPA sample, October 2005 through December 2007 for the post-BAPCPA sample). The points represent the estimated effects of event time (i.e., the $\mu_r$s from the non-parametric event study in equation 5) and the lines represent the parametric event study in equation 6 with the pre-trends normalized between the two periods for ease of visual comparison.

In order to estimate how frequently a hospitalization leads to bankruptcy, we also estimate a parametric event-study specification. This allows us to calculate the "implied effect" at each month relative to hospitalization. We allow for a linear pretrend in event time $m$ (months relative to admission) and a flexible cubic spline with breaks at 0, 12, and 24 months in the post-period. These allow us to estimate the effect of the hospitalization at any point, separately by the hospitalization period:

$$y_{it} = \gamma_t + \mathbb{1}\{\text{Pre-BAPCPA}\}\left(\beta_{0q}m + \beta_{1q}m^2\{m > 0\} + \sum_{s=0}^{2}\beta_{(s+2)q}(m - 12s)^3\{m > 12s\}\right) + \varepsilon_{it}. \quad (6)$$

Figure 6 suggests that the parametric spline fits the non-parametric event-study coefficients well. The identifying assumption requires that, separately for pre-BAPCPA and post-BAPCPA hospitalizations, conditional on having a hospital admission and jointly estimated calendar-year fixed effects, the timing of the admission is uncorrelated with deviations of the outcome from a linear trend in event time.

Table 4: Implied Effects of Hospitalization on Bankruptcy

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Insurance Coverage: | Uninsured | | Insured | |
| Hospitalization Period: | Pre | Post | Pre | Post |
| Implied Effect at 12 Months[a] | 0.89 | 0.18 | 0.19 | 0.15 |
| | (.12) | (.08) | (.08) | (.05) |
| Implied Effect at 24 Months[b] | 1.49 | 0.43 | 0.37 | 0.25 |
| | (0.23) | (0.15) | (0.15) | (0.10) |
| Pre-Hospitalization Mean | 2.11 | 4.94 | 2.07 | 4.38 |
| $p$-value for Null of 12 Month Pre/Post Equality | [<0.001] | | [0.84] | |
| $N$ | 53,611 | 62,912 | 164,207 | 145,502 |

*Notes:* The sample is individuals ages 25-64 who are hospitalized in California, additionally split by the timing of the hospitalization (January 2003 through December 2004 for the pre-BAPCPA sample, October 2005 through December 2007 for the post-BAPCPA sample) and insurance coverage (uninsured or insured which includes those with private insurance or Medicaid coverage). All columns report effects based on OLS estimates of equation 6. The outcome variable is whether an individual has filed for bankruptcy since the beginning of the sample (January 2002). Standard errors (clustered on the individual) are in parentheses. The universe of qualifying uninsured hospitalizations are included in the sample; estimates for the insured are weighted to adjust for individuals' sampling probabilities. All implied effects are significant with p-values less than or equal to .015.

[a] The implied effect at 12 months is calculated from equation 6 as $144 \times \beta_2 + 1,728 \times \beta_4$

[b] The implied effect at 24 months is calculated from equation 6 as $576 \times \beta_2 + 13,824 \times \beta_4$

Figure 6 presents the results of both event studies for the probability an uninsured hospitalization resulted in a bankruptcy filing. The red-diamond markers trace the path of individuals hospitalized in the pre-BAPCPA environment, while the blue-circle markers trace the path of those hospitalized in the post-BAPCPA environment. The pre-BAPCPA hospitalizations result in a pronounced spike in bankruptcy filings following hospitalization, increasing starkly around the time debt is typically sent to collections (around 180 days after the hospitalization). The rate of filings remains persistently higher for the subsequent four years. By comparison, those hospitalized after changes to the bankruptcy code were implemented display a muted filing response to the hospitalization.

Table 4 provides estimates of the "implied effect" of the hospitalization at 12 and 24 months, separately by bankruptcy regime and insurance coverage. The implied effect is the deviation of the parametric coefficients from the linear pretrend, which we interpret as the impact of the hospitalization on whether or not the individual has filed for bankruptcy.

After the reform, uninsured hospitalizations were much less likely to be discharged through bankruptcy. At 24 months post-hospitalization, the pre-BAPCPA uninsured are 1.49 percentage points more likely to file for bankruptcy due to the hospitalization. After implementation, the implied effect of the hospitalization on filing for bankruptcy is just 0.43. While the share of individuals eligible to file for bankruptcy is smaller post-BAPCPA, this is a mechanical result of the construction of a stock variable for ever filing for bankruptcy over the sample period. As further reassurance on this point, the insured demonstrate a similar increase in the share of individuals who have filed for bankruptcy in advance of their hospitalization (4.38 percent from 2.07 percent versus 4.94 percent from 2.11 percent for the uninsured), but a substantially smaller decline after the reform. The marked decline in the implied effect of a hospitalization on filing for bankruptcy indicates that bankruptcy reform significantly reduced the share of uninsured individuals who access bankruptcy as implicit health insurance.

This effect does not appear to be driven by differences in medical debt. Uninsured hospitalizations result in a similar amount of debt sent to collections under both bankruptcy regimes, but 70 percent fewer bankruptcy filings after the reform. Appendix Table A10 shows the implied effect on debt sent to collections 24 months after the hospitalization increased from $6,700 to $6,900 after the reform. While hospitalizations in and of themselves may make up a small share of overall bankruptcy filings (Dobkin et al., 2018b), to the degree that uninsured health shocks can be generalized to other types of uninsured shocks, these results suggest that the reform meaningfully reduced the insurance value of bankruptcy.

## 8 Conclusion

On the one hand, the option to file for bankruptcy provides a form of insurance for American households and lowers the risk of borrowing by providing a process for them to discharge their debts. On the other hand, the option of bankruptcy increases the cost of borrowing, and hence the cost consumers face to smooth consumption over time, by limiting the ability of borrowers to commit to repayment. The existence and relative magnitudes of these two forces have considerable policy implications and are a matter of contentious debate. This paper evaluates changes to the bankruptcy code to demonstrate this trade-off.

Bankruptcy reform induced a net reduction of more than one million bankruptcy filings in the two years after implementation. We demonstrate that this reduction in the risk of bankruptcy filing was passed-through to consumers in the form of lower borrowing costs. The results suggest that a 1-percentage-point reduction in bankruptcy filing risk leads to a 66 basis-point decline in offered credit-card interest rates. Using our model to calibrate a perfectly competitive benchmark, these results imply a pass-through rate of roughly 60 percent. The incomplete pass-through that we find lines up with recent results using other shocks to the consumer credit market (such as Agarwal et al. 2014), and is consistent with some degree of imperfect competition.[39]

Policymakers intended the law's means test to deter higher-income filers, but we find that the income distribution of filers remained essentially unchanged in the wake of the reform. In addition, we find that those hospitalized without health insurance were less likely to declare bankruptcy after the reform. This suggests that BAPCPA decreased the insurance value of bankruptcy.

Collectively, the findings emphasize the trade-offs in determining the optimal generosity of the bankruptcy code. More-generous insurance comes at the cost of higher interest rates. This paper's estimates can inform debates over future changes to the bankruptcy system, changes to other social-insurance programs, and to the regulation of credit markets. To infer the overall

[39]The pass-through results are somewhat larger in magnitude than the estimates reported by Agarwal et al. (2014), but our focus is on interest rates instead of credit limits. A full accounting for the difference in pass-through estimates is outside the scope of this paper, but we speculate that the salience and persistence of the shock that we are studying may account for some of the difference between the estimates. In particular, the BAPCPA reform was widely publicized and well-understood; the huge "rush-to-file" by consumers that we document strongly suggests that other market participants (such as creditors) were likely well-aware of the new law, as well. Additionally, the reform was likely expected to be permanent, which could help creditors overcome their reluctance to modify interest rates in the face of adjustment costs.

welfare impact of the reform would require, at the very least, quantifying the money-metric loss of insurance value for the marginally deterred filers, as well as the gain from lower interest rates for the affected credit-score segments. Strategic borrowing responses to the reform are also difficult to weigh: less generous bankruptcy may decrease moral hazard attributable to the option to discharge debts (Indarte, 2018); but, reducing the generosity of bankruptcy also reduces its salutary effects on entrepreneurship (Fan and White, 2003). We do not make the assumptions or impose the structural framework required to perform this welfare exercise. Nevertheless, the results above identify and quantify a number of the critical inputs for this exercise, which we leave for future research.

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# A Model Derivations and Extensions

For simplicity, we omit subscript $i$ in this section, because the derivations for $i = H, L$ are the same.

## A.1 Interest Rate Comparative Statics

We begin by stating a formal Proposition that summarizes effects of bankruptcy reform on interest rates:

**Proposition 1.** *Under perfect competition, the effect of a change in the exemption level or a change in filing costs on interest rates is given by:*

$$dr/de = \frac{cf(y^*) + (F(y^*) - F(e))}{-bcf(y^*) + b(1-p)},$$
$$dr/dc = \frac{-cf(y^*)}{-bcf(y^*) + b(1-p)},$$

In each case, the sign in the numerator is unambiguous. Both $dr/de$ and $dr/dc$ include $cf(y^*)$, the additional amount of debt discharged rather than repaid by marginal filers who are induced to file by changes in the exemption level or cost of filing. For intuition on this term, recall that filers repay $y - e$ and non-filers repay $(1 + r)b$ in full, and at $y^* = e + (1 + r)b - c$, this difference is $c$. Therefore, $c$ represents the amount that is not repaid to creditors by marginal filers, whose prevalence is represented by $f(y^*)$. Naturally, increases in filing costs and exemptions have opposite effects on the decision to file.

The second term in the numerator of $dr/de$ represents the additional amount discharged rather than paid back to creditors due to changes in repayment behavior for *infra-marginal* filers. If this group is small (for example, because not many filers lie in the mass between $e$ and $y^*$), then the second term becomes less important. Changes in the cost of filing have no effect on the amount recovered by creditors for infra-marginal filers.

Both expressions have the same denominator, which has an ambiguous sign due to the negative first term. It is counter-intuitive that the sign of $dr/dc$ is ambiguous—one would expect a less-generous bankruptcy code to unambiguously lead to lower interest rates. However, there is an additional indirect effect that complicates such a prediction. An increase in $c$ and $e$ changes the decision rule, causing fewer individuals to file for bankruptcy. Thus $r$ increases until the share of individuals filing, $p$, increases to restore $R(r) = b$. The sign is determined by the share of non-filers $(1 - p)$ who repay in full against the additional repayment $c$ from marginal filers.

The derivation of the above result is given in the next subsection.

## A.2 Derivations

We want to derive the following proposition.

**Proposition 2.** *Under perfect competition, the effect of a change in the exemption level or a change in filing costs on interest rates is given by:*

$$\begin{aligned} dr/de &= \frac{cf(y^*) + (F(y^*) - F(e))}{-bcf(y^*) + b(1-p)}, \\ dr/dc &= \frac{-cf(y^*)}{-bcf(y^*) + b(1-p)}. \end{aligned}$$

Recall, with the assumption of perfect competition, we can implicitly define the interest rate $r$ by setting the repayment rate to creditors equal to the amount of borrowing ($R(r) = b$). Observing that $R(r) - b = 0$, we can take partial derivatives in order to apply the implicit function theorem to derive $dr/de$ and $dr/dc$. The direct effects of $e$ and $c$ are straightforward, but the effect of $r$ on $R(r)$ is ambiguous:

$$\frac{\partial R}{\partial c} = \underbrace{cf(y^*)}_{\text{Reduces filings}} > 0$$

$$\frac{\partial R}{\partial e} = \underbrace{-cf(y^*) - F(y^*) + F(e)}_{\text{Increases filings \& reduces recovery (among filers)}} < 0$$

$$\frac{\partial R}{\partial r} = \underbrace{-bcf(y^*)}_{\text{Increases filings}} + \underbrace{b(1-p)}_{\text{Increases recovery among non-filers}} \lesseqgtr 0$$

Using the partial derivatives above, the proposition follows by the implicit function theorem, and

$$\frac{dr}{dc} = -\frac{\partial R/\partial c}{\partial R/\partial r} = \frac{-cf(y^*)}{b(1-p-cf(y^*))}$$

$$\frac{dr}{de} = -\frac{\partial R/\partial e}{\partial R/\partial r} = \frac{cf(y^*) + F(y^*) - F(e)}{b(1-p-cf(y^*))}.$$

We also want to derive the total derivatives for $\frac{dp}{de}$ and $\frac{dp}{dc}$, which we use to derive the empirical object of interest (i.e., $\frac{dr/dc}{dp/dc}$ and $\frac{dr/de}{dp/de}$). To obtain $\frac{dp}{de}$, we can make the following substitutions:

$$\begin{aligned} dp/de &= \partial p/\partial e + \partial p/\partial r * dr/de \\ &= f(y^*) + \frac{dr}{de} bf(y^*) \\ &= f(y^*)\frac{1-F(y^*)}{1-p-cf(y^*)}. \end{aligned}$$

We can do the same for $\frac{dp}{dc}$:

$$\begin{aligned} dp/dc &= \partial p/\partial c + \partial p/\partial r * dr/dc \\ &= -f(y^*) + \frac{dr}{dc} bf(y^*) \\ &= -(f(y^*)\frac{1-F(y^*)}{1-p-cf(y^*)}. \end{aligned}$$

We can use the total derivatives for $\frac{dp}{dc}$, $\frac{dr}{dc}$, $\frac{dp}{de}$, and $\frac{dr}{de}$ to define:

$$\frac{dr/dc}{dp/dc} = \frac{c/b}{1-p}$$

$$\frac{dr/de}{dp/de} = \frac{cf(y^*) + F(y^*) - F(e)}{bf(y^*)(1-F(e))},$$

as desired.

### A.3 Incorporating Insolvency

We can extend the model in Section 3 to incorporate insolvency; that is, the case where we require income of at least $c$ to file bankruptcy so that individuals with income $y < c$ are insolvent and unable to file for bankruptcy.

The filing rule now becomes

$$c \leq y \leq e - c + (1+r)b.$$

The filing probability is now $p = F(e-c+(1+r)b) - F(c) = F(y_h^*) - F(y_l^*)$, where $y_h^*$, $y_l^*$ are upper and lower bounds of filers' income. We assume individuals who cannot afford to file for bankruptcy repay the debt. This reflects wage garnishment or aggressive debt collection. We will assume that whenever $y < c$, individuals repay $y$. As before, individuals with $y > e-c+(1+r)b$ repay $(1+r)b$.

Assuming perfect competition, the equilibrium interest rate is implicitly defined by $R(r) = b$, and the new expression for the expected amount recovered from the population $R(r)$ is

$$R(r) = \underbrace{\int_0^c yf(y)dy}_{\text{Recovered from insolvent}} + \underbrace{\int_e^{e+(1+r)b-c} (y-e)f(y)dy}_{\text{Recovered from bankruptcy filers}} + \underbrace{\int_{e+(1+r)b-c}^{\infty} (1+r)bf(y)dy}_{\text{Recovered from non-filers}}.$$

We can walk through the propositions and empirical object derivations to see how incorporating insolvency changes the expressions. We will find that, while it adds another group of marginal filers, the expressions are qualitatively similar as in the model without insolvency.

**Proposition 3** *The direct effect of a change in the exemption level on probability of filing bankruptcy, and the effect of a change in the cost of filing on probability of filing bankruptcy are given by the following*

$$\begin{aligned}
\partial p/\partial e &= f(e+(1+r)b-c) \\
&= f(y_h^*) > 0, \\
\partial p/\partial c &= -f(e+(1+r)b-c) - f(c) \\
&= -f(y_h^*) - f(y_l^*) < 0.
\end{aligned}$$

The signs are the same as those in the model without insolvent individuals, but a change in the cost of filing now affects two marginal groups: those on the margin of insolvency ($y_l^* = c$); and, the margin in the main model at the asset exemption level: ($y_h^* = e + (1-r)b - c$). An increase in the cost of filing shifts both groups from filing to non-filing.

We can also derive the effects of changes to the bankruptcy code (i.e., $c$, $e$) on interest rates, by re-deriving Proposition 2.

**Proposition 4** *The total effect of a change in exemption level or cost of filing on interest rates are given by the following:*

$$\begin{aligned}
dr/de &= \frac{cf(y_h^*) + F(y_h^*) - F(e)}{b(1 - F(y_h^*) - cf(y_h^*))}, \\
dr/dc &= \frac{-c(f(y_l^*) + f(y_h^*))}{b(1 - F(y_h^*) + cf(y_h^*))}.
\end{aligned}$$

To derive these expressions, first note that $R(r) - b = 0$, then

$$\begin{aligned}
\partial R/\partial r &= b(y-e)f(y)|_{y=e+(1+r)b-c} - b(1+r)bf(y)|_{y=e+(1+r)b-c} + \int_{e+(1+r)b-c}^{\infty} bf(y)dy \\
&= b(1 - F(y_h^*) - cf(y_h^*)), \\
\partial R/\partial c &= yf(y)|_{y=c} - (y-e)f(y)|_{y=e+(1+r)b-c} - (1+r)bf(y)|_{y=e+(1+r)b-c} \\
&= c(f(y_l^*) + f(y_h^*)), \\
\partial R/\partial e &= (y-e)f(y)|_{y=e+(1+r)b-c} - (y-e)f(y)|_{y=e} - (1+r)bf(y)|_{y=e+(1+r)b-c} + \int_{e}^{e+(1+r)b-c} -f(y)dy \\
&= -cf(y_h^*) - F(y_h^*) + F(e).
\end{aligned}$$

We can apply the implicit function theorem to obtain our desired total derivatives:

$$\begin{aligned}
dr/de &= -\frac{\partial R/\partial e}{\partial R/\partial r} \\
&= \frac{cf(y_h^*) + F(y_h^*) - F(e)}{b(1 - F(y_h^*) - cf(y_h^*))}, \\
dr/dc &= -\frac{\partial R/\partial c}{\partial R/\partial r} \\
&= \frac{-c(f(y_l^*) + f(y_h^*))}{b(1 - F(y_h^*) - cf(y_h^*))}.
\end{aligned}$$

The signs and intuition of these total effects are the same as those in the model without insolvent individuals, with additional terms to reflect the filers on the margin of insolvency.

Before we derive the empirical object ($\frac{dr/dc}{dp/dc}$, we again calculate the total derivatives on the filing probability:

$$\begin{aligned}
dp/de &= \partial p/\partial e + \partial p/\partial r * dr/de = f(y_h^*) + \frac{dr}{de} bf(y_h^*) \\
&= f(y_h^*)\frac{1 - F(e)}{1 - F(y_h^*) - cf(y_h^*)} \\
dp/dc &= \partial p/\partial c + \partial p/\partial r * dr/dc = -f(y_h^*) - f(y_l^*) + \frac{dr}{dc} bf(y_h^*) \\
&= -(f(y_h^*) + f(y_l^*))\frac{1 - F(y_h^*)}{1 - F(y_h^*) - cf(y_h^*)}.
\end{aligned}$$

The intuition is similar to cases discussed above. Deriving the empirical objects without approximation,

$$\frac{dr/de}{dp/de} = \frac{cf(y_h^*) + F(y_h^*) - F(e)}{bf(y_h^*)(1 - F(e))}.$$

$$\frac{dr/dc}{dp/dc} = \frac{c/b}{1 - F(y_h^*)}.$$

As before, if we are willing to assume $F(y_h^*) \approx F(e)$, then

$$\frac{dr/de}{dp/de} \approx \frac{cf(y_h^*)}{bf(y_h^*)(1-F(y_h^*))} = \frac{c/b}{1-F(y_h^*)}.$$

# B Data Appendix

## B.1 PACER Bankruptcy Records

Gross et al. (2014) contacted every bankruptcy court in the US and requested a waiver of PACER fees; 81 districts granted the research team a waiver. They downloaded the dockets for each court from the 1990s through 2011.

For the purposes of this paper, we validated that dataset by comparing the annual counts of bankruptcies to administrative records. We discarded three districts if their annual counts scraped from the PACER database diverged from the official administrative record by more than 10% in any year between 2004 and 2007.[40] The final sample consists of 78 districts over that time period.

## B.2 Consumer Financial Production Bureau Consumer Credit Panel (CCP) and Mintel Data

The Consumer Financial Protection Bureau Consumer Credit Panel (CCP) is a 1-in-48 random sample of U.S. consumers with credit records. Our primary use of the CCP is to estimate the bankruptcy filing risk for each credit score segment. To do so, we combine all public record snapshots in the CCP. We eliminate any duplicate public records to obtain a clean index file, which we merge with the full credit score archives for consumers. Consumers without a credit score are dropped. For each consumer, we allot them to a credit score segment (defined as the 10-point credit score bins). The small number of consumers with credit scores below 440 are allocated to that bin. At each point in time, we estimate the share of the consumers in that credit score segment who file for bankruptcy over the subsequent 12 months. Individuals without credit scores are dropped from the sample.

Data on credit card offers are from Mintel Comperemedia, accessed through the Consumer Financial Protection Bureau. Mintel Comperemedia conducts proprietary market research by surveying United States households, who forward all incoming marketing mail. We focus on credit card offers. The data include rich information on each credit card offer, including card categories (Affinity Cards, Co-Branded, Credit Cards, Lifestyle Cards, Retail Cards, Secured Cards), application type (Confirmed, General, Guaranteed Approval, Pre-Approved, Pre-Qualified, Pre-Selected), and the lender. They additionally include information on the offered interest rate, and whether (and for how long) an introductory ("teaser") rate might be applied. Importantly for our purposes, the offers are coupled with information on the consumer who received the offer, including their credit score and state of residence. We drop offers associated with consumers who are missing credit scores and offers for which interest rates are missing. The data is a repeated cross-section, surveying around 2,500 individuals each month and include between 5,900 and 12,079 credit card offers over our sample period (with both the mean and median number of offers around 8,000 per month).

## B.3 Hospitalizations Data

We use hospital discharge data from the California Office of Statewide Health Planning and Development (OSHPD). The hospitalizations data are merged with credit reports and vital records

[40] Those three districts were MOE, MTB, and NYN.

using social security numbers as described in the Online Appendix of Dobkin et al. (2018a). All data production and analysis happened on-site at OSHPD's Sacramento office and all output was reviewed by OSHPD staff to confirm privacy was protected.

The hospital discharge data includes a unique identifier, dates of admission and discharge, details about the health event (e.g., diagnosis codes), and demographic information. It also includes an indicator for insurance coverage which includes Medicaid, private insurance, and "self-pay." We use the primary payer of the index admission to define insurance coverage.

We sample non-pregnancy related admissions with a non-missing social security number from 2003 through 2007. We additionally use hospitalizations from 2000 to 2010 to limit the sample to admissions which were the first in three years for the individual, in order to isolate health "shocks." We select the universe of "self-pay" (uninsured) hospitalizations. For those insured with Medicaid or private coverage (insured), we sample a random 20% of individuals whose admission originated through the Emergency Department, and a random 10% of individuals whose admission was not through the Emergency Department. We construct reweights according to the inverse probability an individual was sampled. We restrict to ages 25 to 64. For additional sample selection and summary statistics, see Dobkin et al. (2018a).

We convert the credit report variable for bankruptcy filings from a flow into a stock by defining a cumulative indicator variable based on whether the individual has filed for bankruptcy since entering the sample in 2002. This allows the event study specification to exploit variation in the timing of the hospitalization to identify the effect of the hospitalization on the likelihood of filing for bankruptcy.

Finally, we define whether hospitalizations were exposed to the "pre-BAPCPA" or "post-BAPCPA" bankruptcy regime. We define those hospitalized between January 2003 through December 2004 as facing the pre-BAPCPA bankruptcy code and hospitalizations between October 2005 through December 2007 for the post-BAPCPA sample. Most hospitalization-induced bankruptcies occur in the first 18 months following the hospitalization. In order to limit the impact of intertemporally substituted bankruptcies filed during the rush-to-file period just before BAPCPA went into effect, we limit the pre-BAPCPA sample to those hospitalized by the end of 2004. Any individuals hospitalized in or after October 2005 faced the post-BAPCPA bankruptcy code.

# C Appendix Tables

Table A1: Percentage of Total Filings Covered by PACER Sample

| | | (1) | (2) | (3) |
|---|---|---|---|---|
| Year | Quarter | All Bankruptcy Filings | Chapter 7 | Chapter 13 |
| 2004 | 1 | 86.2 | 89.6 | 74.1 |
| | 2 | 85.5 | 89.2 | 74.2 |
| | 3 | 86.1 | 90.6 | 74.5 |
| | 4 | 86.0 | 90.6 | 74.1 |
| 2005 | 1 | 86.1 | 89.8 | 74.4 |
| | 2 | 86.2 | 89.3 | 74.4 |
| | 3 | 87.2 | 90.1 | 75.1 |
| | 4 | 88.1 | 90.0 | 75.4 |
| 2006 | 1 | 82.8 | 87.6 | 74.8 |
| | 2 | 83.1 | 87.5 | 75.0 |
| | 3 | 83.6 | 89.0 | 74.9 |
| | 4 | 84.5 | 91.0 | 74.8 |
| 2007 | 1 | 85.5 | 91.2 | 75.3 |
| | 2 | 86.4 | 92.1 | 75.7 |
| | 3 | 86.3 | 92.8 | 75.4 |
| | 4 | 86.6 | 93.1 | 75.5 |

*Notes:* The table presents the percent of the total administrative counts of bankruptcies which are included in the PACER sample in each year and quarter of the data. Administrative counts are provided by the Administrative Office of the United States Courts.

Table A2: Net Change in Filings through 2007 (Robustness to Counterfactual Specifications)

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | | Panel A. Main Sample Period 2004-2007 | | | | |
| Total | -1,077,679 | -1,085,106 | -1,549,639 | -1,529,728 | -1,637,479 | -1,618,761 |
| Chapter 7 | -946,148 | -948,801 | -1,444,828 | -1,419,240 | -1,533,578 | -1,509,383 |
| Chapter 13 | -160,950 | -173,816 | -157,714 | -158,298 | -153,844 | -154,385 |
| | | Panel B. Extended Sample Period 2002-2007 | | | | |
| Total | -1,109,094 | -859,433 | -1,057,545 | -1,215,139 | -1,244,948 | -1,105,452 |
| Chapter 7 | -873,627 | -687,479 | -996,780 | -900,218 | -1,029,042 | -949,787 |
| Chapter 13 | -231,271 | -173,914 | -218,421 | -164,546 | -215,956 | -162,645 |
| Date Used | Senate | Senate | House | House | Signed | Signed |
| Unemployment Rate | | ✓ | | ✓ | | ✓ |

*Notes:* This table presents robustness to results presented in Table 2. In each column, we estimate the total deviation from the predicted number of bankruptcy filings through the end of 2007. We estimate equation 2 from the beginning of the sample until BAPCPA until the date indicated in the "Date Used" row. The Senate passage date is March 10, 2005, the House passage date is April 14, 2005, and the date signed is April 20, 2005. We additionally include the national unemployment rate in estimating equation 2 where indicated. The overall numbers are inflated to reflect the nation as a whole, based on our PACER sample coverage (see Appendix Table A1).

Table A3: Benchmarking Interest Rate Pass-through

| (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|
| | | | | | | Pass-through Calibration | |
| Credit Score Segment | Population Share | Bankruptcy Rate | APR | APR w/ Fees | 60-Day Default Rate | APR | APR w/ Fees |
| ≤500 | 6.5% | 4.1% | 12.5% | 24.0% | 55.7% | 0.52 | 0.57 |
| 510 | 1.5% | 4.3% | 12.2% | 22.6% | 44.4% | 0.65 | 0.71 |
| 520 | 1.5% | 4.2% | 13.3% | 24.7% | 40.4% | 0.71 | 0.78 |
| 530 | 1.5% | 4.1% | 13.8% | 24.5% | 37.0% | 0.75 | 0.82 |
| 540 | 1.5% | 3.7% | 14.0% | 24.9% | 32.4% | 0.80 | 0.88 |
| 550 | 1.6% | 3.4% | 13.9% | 23.6% | 29.0% | 0.84 | 0.91 |
| 560 | 1.6% | 3.0% | 13.9% | 23.7% | 25.4% | 0.88 | 0.95 |
| 570 | 1.6% | 2.5% | 13.9% | 22.7% | 21.3% | 0.92 | 0.99 |
| 580 | 1.7% | 2.3% | 13.8% | 22.6% | 17.5% | 0.96 | 1.03 |
| 590 | 1.7% | 1.9% | 13.9% | 21.7% | 14.8% | 0.99 | 1.06 |
| 600 | 1.7% | 1.8% | 14.4% | 22.5% | 12.2% | 1.02 | 1.10 |
| 610 | 1.8% | 1.6% | 14.1% | 21.0% | 9.9% | 1.04 | 1.11 |
| 620 | 2.0% | 1.5% | 13.7% | 20.4% | 8.0% | 1.06 | 1.12 |
| 630 | 2.0% | 1.3% | 13.4% | 19.0% | 6.2% | 1.08 | 1.13 |
| 640 | 2.2% | 1.2% | 13.2% | 18.7% | 5.1% | 1.09 | 1.14 |
| 650 | 2.4% | 1.1% | 13.0% | 17.4% | 3.9% | 1.10 | 1.14 |
| 660 | 2.6% | 1.0% | 12.3% | 16.5% | 3.2% | 1.10 | 1.14 |
| 670 | 2.6% | 0.9% | 11.6% | 14.7% | 2.7% | 1.10 | 1.13 |
| 680 | 2.8% | 0.8% | 11.1% | 14.0% | 2.1% | 1.10 | 1.13 |
| 690 | 2.9% | 0.6% | 10.8% | 13.6% | 1.6% | 1.10 | 1.13 |
| 700 | 3.4% | 0.5% | 10.3% | 12.9% | 1.3% | 1.09 | 1.12 |
| 710 | 3.6% | 0.3% | 10.1% | 12.5% | 1.0% | 1.09 | 1.12 |
| 720 | 3.6% | 0.3% | 9.8% | 12.1% | 0.7% | 1.09 | 1.12 |
| 730 | 3.8% | 0.2% | 9.7% | 12.0% | 0.6% | 1.09 | 1.12 |
| 740 | 4.2% | 0.1% | 9.5% | 11.7% | 0.5% | 1.09 | 1.11 |
| 750 | 3.9% | 0.1% | 9.5% | 11.6% | 0.4% | 1.09 | 1.11 |
| 760 | 4.1% | 0.1% | 9.4% | 11.5% | 0.3% | 1.09 | 1.11 |
| 770 | 4.2% | 0.0% | 9.4% | 11.4% | 0.2% | 1.09 | 1.11 |
| 780 | 4.9% | 0.0% | 9.3% | 11.3% | 0.2% | 1.09 | 1.11 |
| 790 | 4.7% | 0.0% | 9.4% | 11.4% | 0.2% | 1.09 | 1.11 |
| 800 | 4.9% | 0.0% | 9.4% | 11.5% | 0.1% | 1.09 | 1.11 |
| 810 | 4.6% | 0.0% | 9.5% | 11.5% | 0.1% | 1.09 | 1.11 |
| 820 | 3.3% | 0.0% | 9.5% | 11.5% | 0.1% | 1.09 | 1.11 |
| 830 | 1.9% | 0.0% | 9.4% | 11.4% | 0.1% | 1.09 | 1.11 |
| 840 | 1.2% | 0.0% | 9.6% | 11.6% | 0.0% | 1.10 | 1.12 |
| **Weighted averages** | | **1.05%** | **11.03%** | **15.46%** | **9.06%** | **1.016** | **1.052** |

*Notes:* This table reports calibration estimates of pass-through estimates for each credit score segment. See main text for details. The lowest credit score bin combines all credit scores below 500. The "Interest rate" column comes from the credit card offers data, and the bankruptcy rate comes from the Consumer Credit Panel (CCP). The pass-through estimate comes from combining the estimates in columns according to equation (1), using default rate as proxy for one minus the recovery rate, which is the first term in equation (1).

Table A4: Summary Statistics for Credit Card Offers

| | (1) | (2) | (3) |
|---|---|---|---|
| | Prime | Subprime | All Borrowers |
| APR | 11.50 | 14.52 | 11.88 |
| Weighted APR | 6.61 | 10.67 | 7.12 |
| Introductory APR | 5.46 | 8.76 | 5.87 |
| Rate Spread | 4.85 | 7.58 | 5.19 |
| Weighted Rate Spread | -0.04 | 3.73 | 0.43 |
| | | | |
| Pre-Approved | 61.6% | 74.1% | 63.2% |
| Annual Fee | 11.0% | 52.5% | 16.2% |
| Rewards | 59.5% | 16.7% | 54.1% |
| Annual Fee, No Rewards | 4.6% | 49.8% | 10.2% |
| Introductory Rate | 56.3% | 43.6% | 54.7% |
| | | | |
| Credit Score | 750 | 566 | 727 |
| Mean Offers Per Month | 3.33 | 2.77 | 3.26 |
| $N$ (Individual-Months) | 105,941 | 13,982 | 119,923 |
| $N$ (Offers) | 352,589 | 38,690 | 391,279 |

*Notes:* The sample is individuals in the Mintel sample at any point between January 2004 and December 2007, collapsed to the individual-month. The table presents mean features of credit card offers, weighted by the mail volume of the campaign.

Table A5: Defining $\delta_b$: Change in Prospective Filing Risk

| (1) | (2) | (3) | (4) |
|---|---|---|---|
| Credit Score | Pre-BAPCPA | Post-BAPCPA | $\delta_b$ |
| | Subprime | | |
| 440 | 0.0554 | 0.0193 | -0.0361 |
| 450 | 0.0459 | 0.0170 | -0.0289 |
| 460 | 0.0412 | 0.0154 | -0.0259 |
| 470 | 0.0393 | 0.0156 | -0.0236 |
| 480 | 0.0392 | 0.0162 | -0.0229 |
| 490 | 0.0402 | 0.0179 | -0.0224 |
| 500 | 0.0425 | 0.0189 | -0.0235 |
| 510 | 0.0442 | 0.0206 | -0.0236 |
| 520 | 0.0431 | 0.0212 | -0.0218 |
| 530 | 0.0418 | 0.0201 | -0.0217 |
| 540 | 0.0373 | 0.0183 | -0.0191 |
| 550 | 0.0341 | 0.0167 | -0.0174 |
| 560 | 0.0294 | 0.0149 | -0.0145 |
| 570 | 0.0252 | 0.0127 | -0.0125 |
| 580 | 0.0223 | 0.0115 | -0.0109 |
| 590 | 0.0186 | 0.0101 | -0.0086 |
| 600 | 0.0167 | 0.0089 | -0.0078 |
| 610 | 0.0151 | 0.0081 | -0.0070 |
| 620 | 0.0132 | 0.0075 | -0.0057 |
| | Prime | | |
| 630 | 0.0123 | 0.0069 | -0.0055 |
| 640 | 0.0112 | 0.0064 | -0.0049 |
| 650 | 0.0105 | 0.0058 | -0.0047 |
| 660 | 0.0091 | 0.0051 | -0.0040 |
| 670 | 0.0085 | 0.0048 | -0.0037 |
| 680 | 0.0070 | 0.0038 | -0.0032 |
| 690 | 0.0060 | 0.0034 | -0.0026 |
| 700 | 0.0043 | 0.0025 | -0.0018 |
| 710 | 0.0032 | 0.0018 | -0.0014 |
| 720 | 0.0025 | 0.0015 | -0.0010 |
| 730 | 0.0017 | 0.0010 | -0.0007 |
| 740 | 0.0011 | 0.0007 | -0.0004 |
| 750 | 0.0010 | 0.0006 | -0.0004 |
| 760 | 0.0007 | 0.0005 | -0.0002 |
| 770 | 0.0005 | 0.0003 | -0.0001 |
| 780 | 0.0003 | 0.0002 | -0.0001 |
| 790 | 0.0002 | 0.0001 | -0.0001 |
| 800 | 0.0001 | 0.0001 | $<.0001$ |
| 810+ | $<.0001$ | $<.0001$ | $<.0001$ |

*Notes:* The sample is individuals with a non-missing credit score in the CFPB CCP from September 2003 through December 2007. Columns 2 and 3 present the average 12-month prospective bankruptcy filing probabilities before and after bankruptcy reform, respectively. Column 4 presents the difference.

Table A6: Pass-through to Interest Rates (using pre-BAPCPA risk)

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Dependent variable: | Regular Interest Rate | | Adjusted Interest Rate | |
| Post-BAPCPA $\times \delta_b$ | -62.1*** | -44.1** | -37.9** | -28.5* |
| | (19.68) | (17.0) | (18.0) | (15.1) |
| Baseline Controls | ✓ | ✓ | ✓ | ✓ |
| Subprime$\times t$ | | ✓ | | ✓ |
| $N$ | 390,975 | 390,975 | 390,975 | 390,975 |

*Notes:* The sample is credit card offers made to households from January 2004 through December 2007. All columns report effects based on OLS estimates of equation 4, with the treatment $\delta_b$ replaced by the June 2004 prospective 12-month filing probability for the credit score bin $b$. The outcome variables are the interest rates on credit card offers, adjusted for the prime rate. Baseline controls include month-by-year and lender fixed effects. Standard errors (two-way clustered by credit score segment and lender) are in parentheses. Offers are weighted by the mail volume of the campaign. Asterisks indicate significance at the 1 percent (***), 5 percent (**), and 10 percent (*) level, respectively.

Table A7: Pass-through to Adjusted Interest Rates (Robustness)

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Dependent variable: | | | | | | | | |
| Post-BAPCPA $\times \delta_b$ | -29.74 | -12.8 | -73.1** | -36.1 | -32.8 | -36.1 | -36.5 | -35.9 |
| | (61.9) | (61.3) | (32.9) | (29.7) | (24.5) | (25.1) | (25.2) | (24.8) |
| $R^2$ | 0.08 | 0.10 | 0.36 | 0.36 | 0.44 | 0.45 | 0.45 | 0.45 |
| Month $\times$ Year FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Score Bin | | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Lender FEs | | | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Subprime $\times t$ | | | | ✓ | ✓ | ✓ | ✓ | ✓ |
| Card Category | | | | | ✓ | ✓ | ✓ | ✓ |
| Application Type | | | | | | ✓ | ✓ | ✓ |
| State FE | | | | | | | ✓ | ✓ |
| Census Division $\times$ Year FE | | | | | | | | ✓ |
| $N$ | 391,279 | 391,153 | 391,153 | 391,153 | 390,975 | 390,975 | 390,975 | 390,381 |

*Notes:* The sample is credit card offers made to households from January 2004 through December 2007. All columns report effects based on OLS estimates of equation 4. The outcome variable is the adjusted interest rate on credit card offers, adjusted for introductory interest rates. Standard errors (two-way clustered by credit score segment and lender) are in parentheses. Offers are weighted by the mail volume of the campaign. Asterisks indicate significance at the 1 percent (***), 5 percent (**), and 10 percent (*) level, respectively.

Table A8: Comparing Difference-in-Difference Results to Heterogeneous Treatment Effects Model

| | (1) | (2) |
|---|---|---|
| Dependent variable: | Regular Interest Rate | |
| | Main result, Table 3, column (4) | Benchmark DiD estimate |
| Post-BAPCPA $\times \delta_b$ | -60.35** | |
| | (27.18) | |
| Post-BAPCPA $\times$ (FICO<650) | | -0.825*** |
| | | (0.148) |
| Implied interaction term (to compare to Table 3, column (4)) | | -51.98*** |
| Analysis of Average Treatment Effect (ATE) estimate | | |
| # of negative weights in calculation of ATE | | 0 |
| Standard deviation of weights in ATE calculation | | 0.0007 |
| Ratio of DiD coeffcient to standard deviation of weights | | 1,147 |
| $R^2$ | 0.433 | 0.689 |
| Month-x-Year FE | ✓ | ✓ |
| Score Bin | ✓ | ✓ |
| Lender FEs | ✓ | ✓ |
| Subprime $\times t$ | ✓ | ✓ |
| N | 391,279 | 1,948 |

*Notes:* Column (1) reproduces the main result in Table 3, column (4). See Table 3 for more details on specification. Column (2) reports several different results based on a "collapsed" data set that collapses the micro data by year-month and credit score "bin", where each of these indicator variables (as well as subprime-specific trend) has been residualized by lender fixed effects. The first row in column (2) reports a benchmark difference-in-difference (DiD) coefficient estimate using analogous specification to column (1), but replacing the continuous treatment intensity ($\delta_b$) with an indicator variable for whether FICO score is below 650 (which is close to median FICO score in the sample). This results in a standard DiD estimate, and the standard errors are clustered by credit score bin. By scaling this estimate by the difference in average value of $\delta_b$ in sample below and above FICO=650, this scaled estimate corresponds to an implied interaction term which can then be compared to main result in Table 3, column (4). The next 3 rows report statistics based on the de Chaisemartin and D'Haultfoeuille (2019) heterogeneous treatment effects model. The first row reports the number of negative weights used in calculation of ATE. If there are negative weights, then this suggests reporting DiD estimates that are robust to treatment effect heterogeneity. The lack of negative weights and very small standard deviation of weights suggests that the benchmark ATE is likely to be reliable, which is also similar to the main results estimated on the full micro data.

Table A9: Summary Statistics for Hospitalizations

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Insurance Coverage: | Uninsured | | Insured | |
| Hospitalization Period: | Pre | Post | Pre | Post |
| Age | 44 | 45 | 48 | 49 |
| | (11) | (11) | (10) | (10) |
| Asian | 0.046 | 0.046 | 0.066 | 0.07 |
| | (.21) | (.21) | (.25) | (.25) |
| Black | 0.11 | 0.11 | 0.077 | 0.077 |
| | (.31) | (.31) | (.27) | (.27) |
| Hispanic | 0.24 | 0.27 | .18 | .19 |
| | (.43) | (.44) | (.38) | (.39) |
| White | 0.56 | 0.53 | .64 | .62 |
| | (.5) | (.5) | (.48) | (.48) |
| Male | 0.62 | 0.61 | .45 | .45 |
| | (.49) | (.49) | (.5) | (.5) |
| Chronic Diagnosis | 0.78 | 0.82 | 0.84 | 0.87 |
| | (.41) | (.38) | (.36) | (.34) |
| Zip Code Median Income | 59,146 | 58,957 | 66,652 | 67,307 |
| | (22,013) | (21,866) | (24,307) | (24,505) |
| Any Collection in Last 12 Months | 0.34 | 0.38 | .16 | .17 |
| | (.47) | (.49) | (.36) | (.38) |
| Collection Balance | 2,869 | 3,994 | 1,068 | 1,341 |
| | (9,181) | (11,528) | (5,834) | (6,481) |
| Any Bankruptcy in Last 12 Months | 0.014 | 0.012 | .014 | .011 |
| | (.12) | (.11) | (.12) | (.1) |
| Credit Limit | 13,366 | 16,368 | 30,164 | 43,741 |
| | (39,116) | (51,555) | (51,750) | (80,580) |
| Credit Score | 655 | 655 | 727 | 734 |
| | (111) | (109) | (119) | (120) |
| *N* | 53,611 | 62,912 | 164,207 | 145,502 |

*Notes:* The sample is individuals ages 25-64 who are hospitalized in California, additionally split by the timing of the hospitalization (January 2003 through December 2004 for the pre-BAPCPA sample, October 2005 through December 2007 for the post-BAPCPA sample) and insurance coverage (uninsured or insured which includes those with private insurance or Medicaid coverage). Age is defined at admission. Insurance status is defined at the index admission and denotes coverage by Medicaid or private insurance. The universe of qualifying uninsured hospitalizations are included in the sample; estimates for the insured are weighted to adjust for individuals' sampling probabilities. Standard deviations are in parentheses.

Table A10: Implied Effects of Hospitalization on Debt in Collections

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Insurance Coverage: | Uninsured | | Insured | |
| Hospitalization Period: | Pre | Post | Pre | Post |
| | *Panel A. Number of Debts in Collections* | | | |
| Implied Effect at 12 Months | 0.96 | 1.14 | 0.11 | 0.13 |
| | (0.02) | (0.03) | (0.01) | (0.01) |
| Implied Effect at 24 Months | 1.32 | 1.52 | 0.16 | 0.21 |
| | (0.05) | (0.06) | (0.02) | (0.02) |
| Pre-Hospitalization Mean | 1.05 | 3.06 | 0.43 | 1.28 |
| | *Panel B. Collection Balance* | | | |
| Implied Effect at 12 Months | 4,559 | 5,163 | 103 | 178 |
| | (104.7) | (135.1) | (28.0) | (28.6) |
| Implied Effect at 24 Months | 6,724 | 6,944 | 163 | 316 |
| | (169.0) | (229.1) | (53.1) | (56.8) |
| Pre-Hospitalization Mean | 2,869 | 3,994 | 1,068 | 1,341 |
| $N$ | 53,611 | 62,912 | 164,207 | 145,502 |

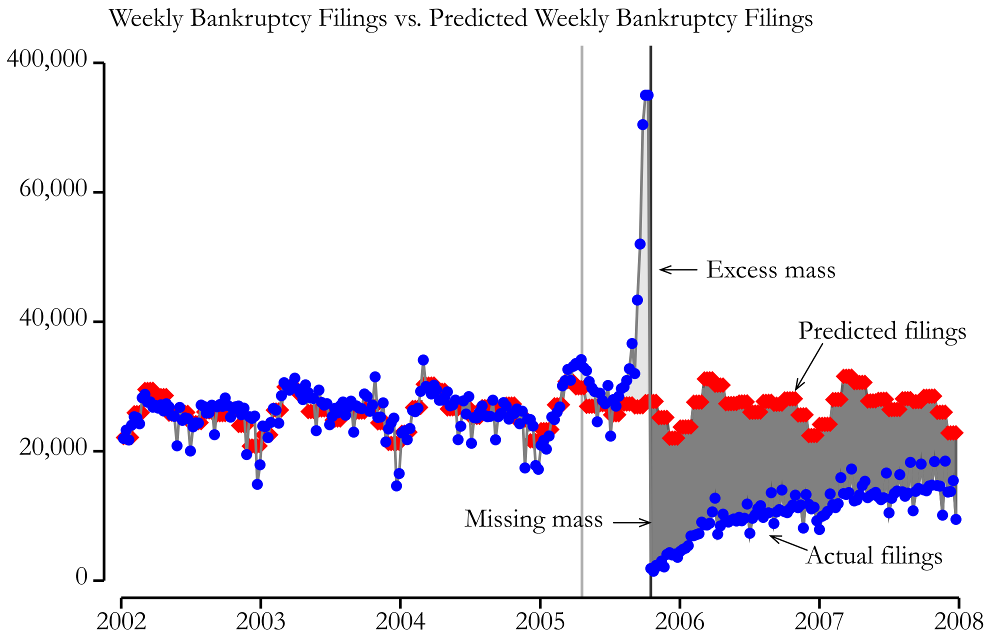
*Notes:* The sample is individuals ages 25-64 who are hospitalized in California, additionally split by the timing of the hospitalization (January 2003 through December 2004 for the pre-BAPCPA sample, October 2005 through December 2007 for the post-BAPCPA sample) and insurance coverage (uninsured or insured which includes those with private insurance or Medicaid coverage). All columns report effects based on OLS estimates of equation 6. The outcome variable is whether an individual has filed for bankruptcy since the beginning of the sample (January 2002). Standard errors (clustered on the individual) are in parentheses. The universe of qualifying uninsured hospitalizations are included in the sample; estimates for the insured are weighted to adjust for individuals' sampling probabilities. All implied effects are significant at the 1% level.

[a] The implied effect at 12 months is calculated from equation 6 as $144 \times \beta_2 + 1,728 \times \beta_4$

[b] The implied effect at 24 months is calculated from equation 6 as $576 \times \beta_2 + 13,824 \times \beta_4$

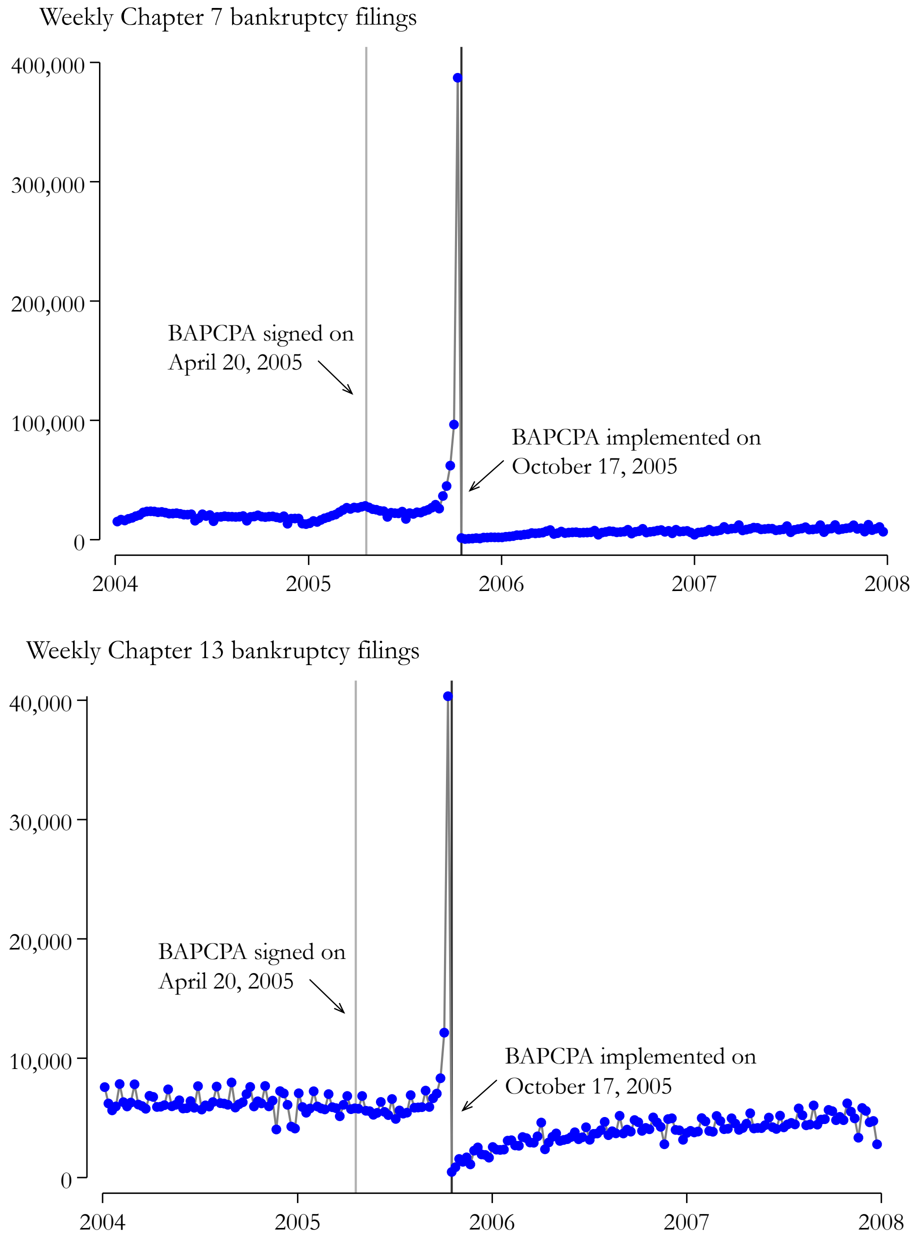
# D Appendix Figures

Figure A1: Excess and Missing Mass of Bankruptcy Filings: Extended Pre-Period



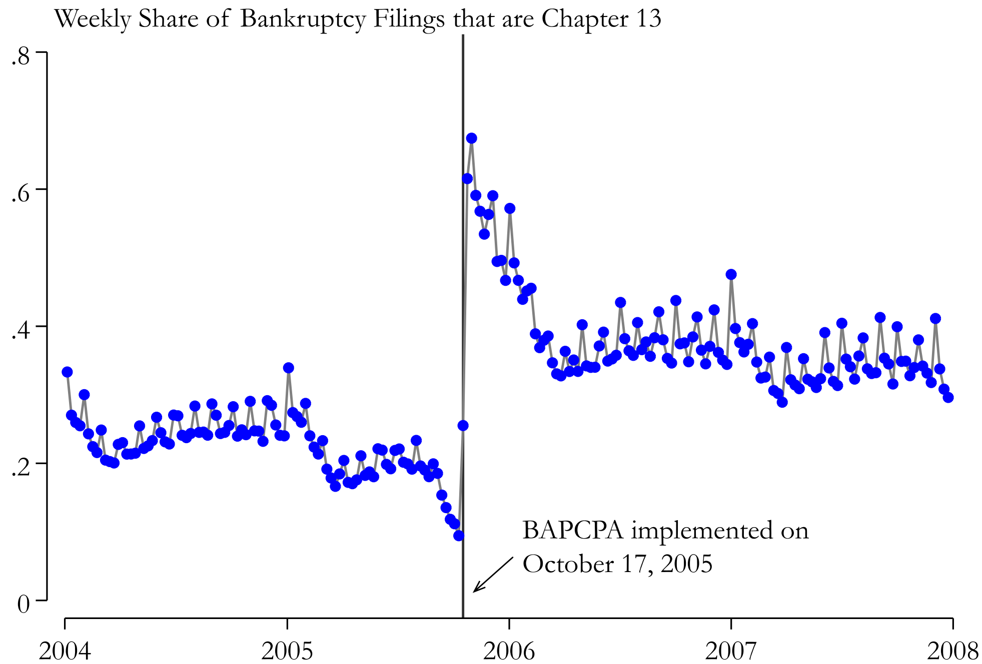

*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2002 through December 2007. The total count of filings for each week is plotted against the predicted number of filings for the week. The predicted number of filings are the result of estimating equation 2 on the total count of filings from January 2004 through the day that BAPCPA was passed by the Senate (March 10, 2005). The two data points before implementation of BAPCPA are censored in this figure: there were 108,745 filings during the week that began on October 3, 2005 and 427,947 filings during the week that began on October 10, 2005.

Figure A2: Time Series for Chapter 7 and Chapter 13 Filings



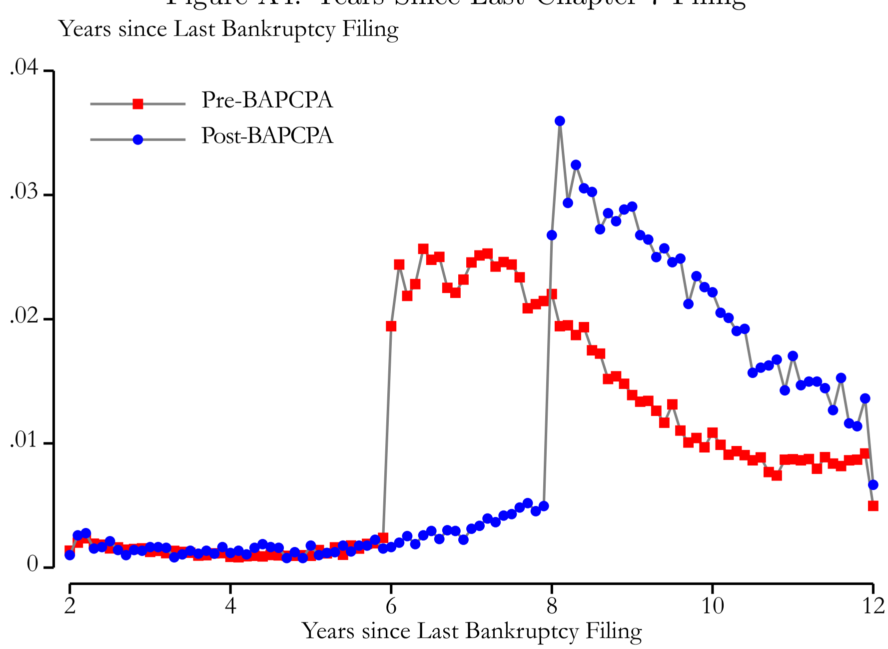

*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007. Each dot in the figure represents the total count of filings for that week, separately for Chapter 7 filings (top figure) and Chapter 13 filings (bottom figure).

Figure A3: Share Chapter 13



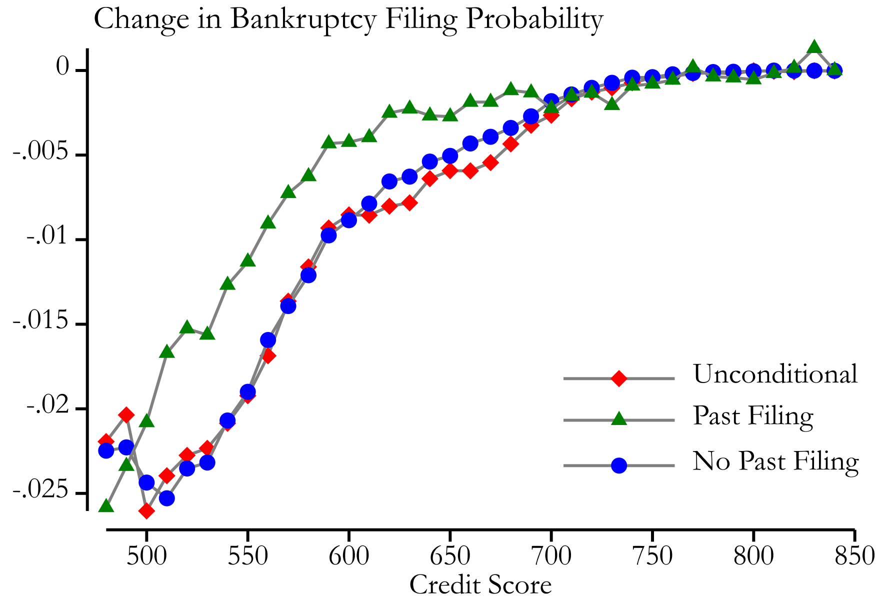

*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007. Each dot in the figure represents the share of filings in that week which were Chapter 13. The vertical line indicates the date when BAPCPA was implemented, October 17, 2005.

Figure A4: Years Since Last Chapter 7 Filing



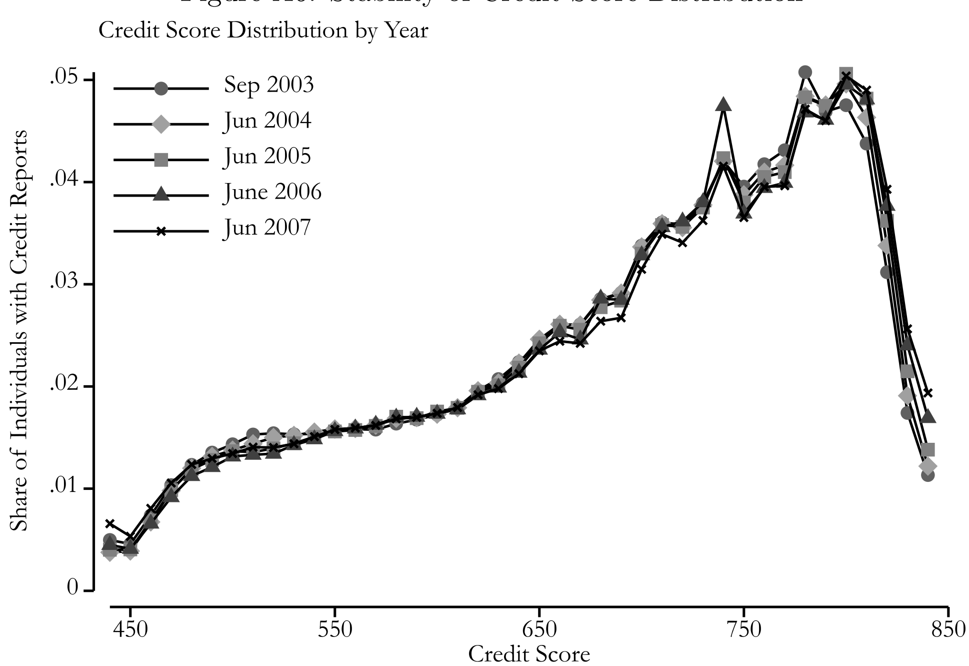

*Notes:* The sample includes Chapter 7 consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007. We match filings for the same individual over time using name, last four digits of Social Security number, and district. The figure plots the distribution of "years since last Chapter 7 filing" for bankruptcies filed before and after BAPCPA was implemented (October 17, 2005).

Figure A5: Change in Probability of Filing by Filing Status



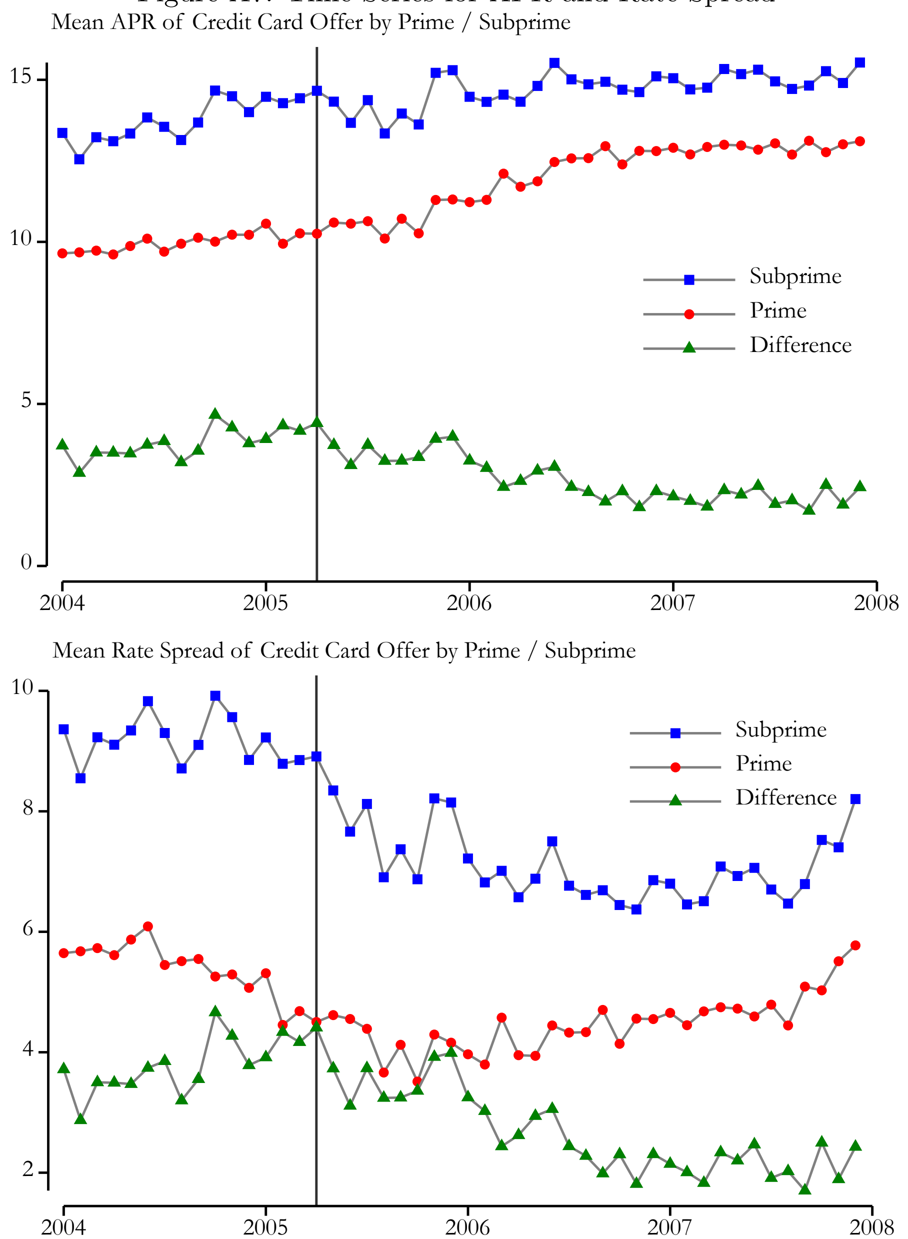

*Notes:* The sample are individuals in the CFPB CCP. This figure presents the change in the probability of individuals in a 10-point credit score segment who file for bankruptcy within the next 12 months (before and after the implementation of BAPCPA). Each point represents the change in the filing rate for a 10-point credit score segment, additionally split by whether the individuals have a prior bankruptcy record dating back to the first CCP observation in 2001.

Figure A6: Stability of Credit Score Distribution



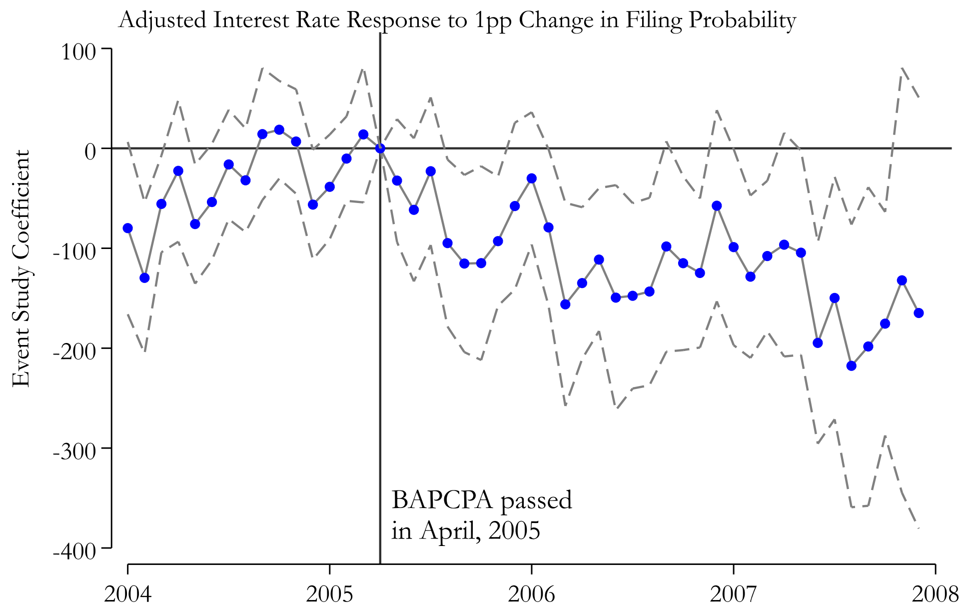

*Notes:* The sample is individuals with a non-missing credit score in the CFPB CCP from September 2003 through December 2007. The points represent the share of consumers with a credit score in the 10-point credit score bin.

Figure A7: Time Series for APR and Rate Spread



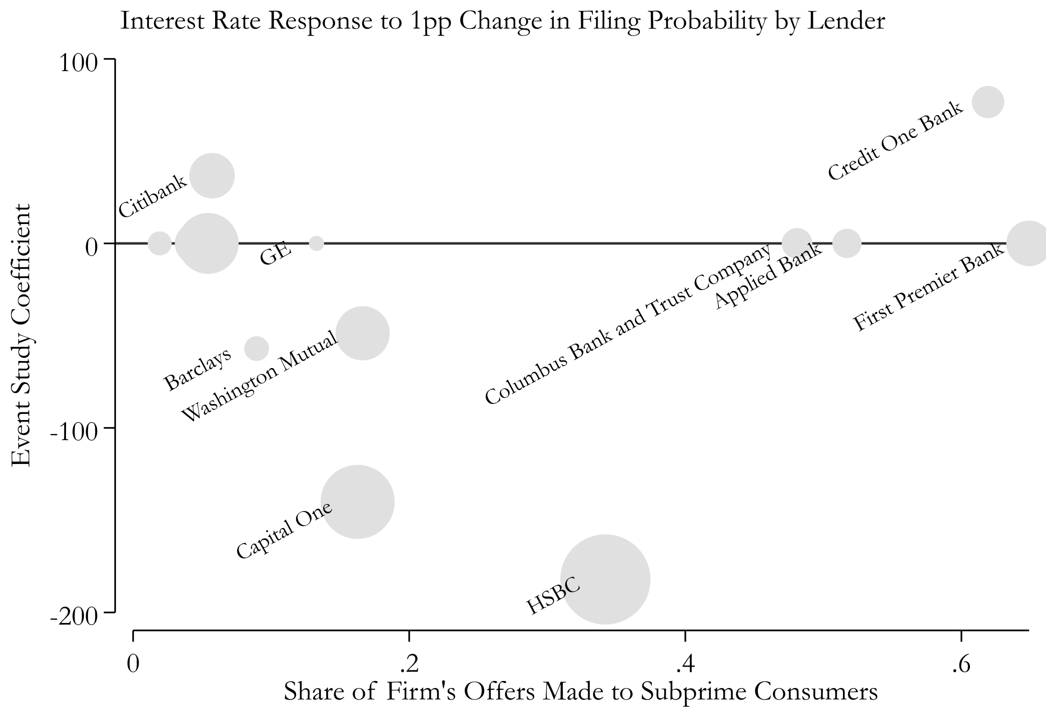

*Notes:* The sample is credit card offers made between January 2004 and December 2007 included in the Mintel data. Figures plot the average interest rate (either APR or rate spread) offer made to prime and subprime borrowers (defined as a credit score 620 or below).

Figure A8: Effect of Decline in Filing Probability on Offered Interest Rates



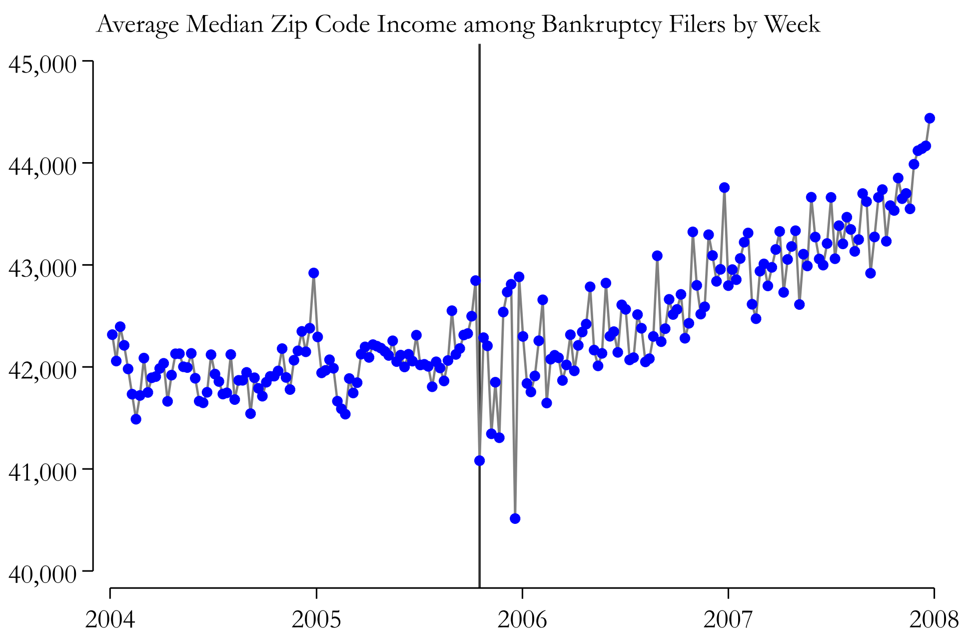

*Notes:* The sample is credit card offers made between January 2004 and December 2007 included in the Mintel data. The points represent estimates of the $\lambda_t$'s in equation 3. The dashed lines provide 95% confidence intervals for each point. The dependent variable is the rate spread for offered interest rate adjusted for the teaser rate.

Figure A9: Lender-Specific Interest Rate Responses to Change in Filing Rates



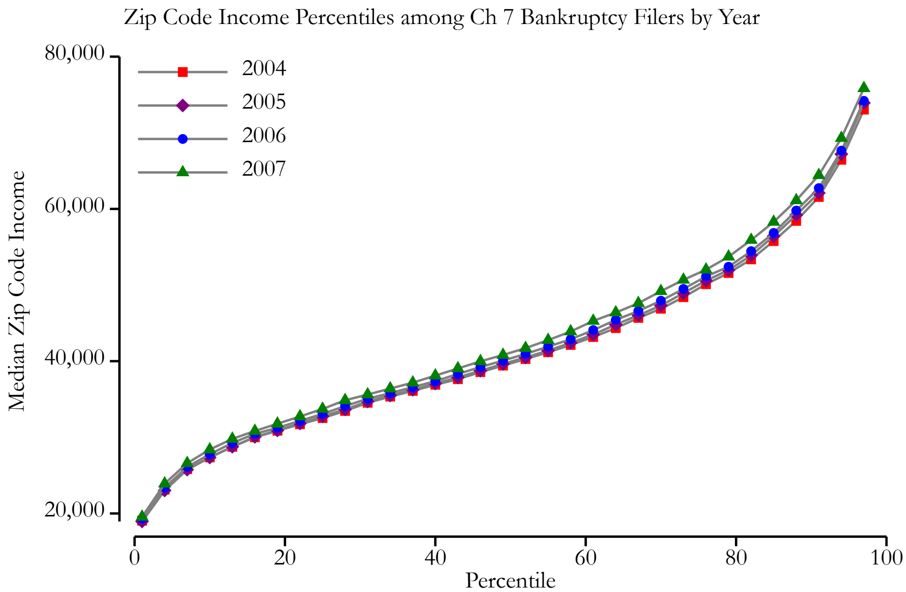

*Notes:* The sample is credit card offers made between January 2004 and December 2007 included in the Mintel data. The figure plots the coefficient $\beta_1$ from the estimation of equation 4 separately by each lender, which denotes the change in the interest rates on credit card offers for a 1 percentage point change in the probability an individual in a given credit score segment files for bankruptcy. The x-axis denotes the share of credit card offers by a firm which are made to subprime consumers, while the size of the circle is determined by the total number of subprime offers made during the sample period. All offers are weighted by the mail volume of the campaign.

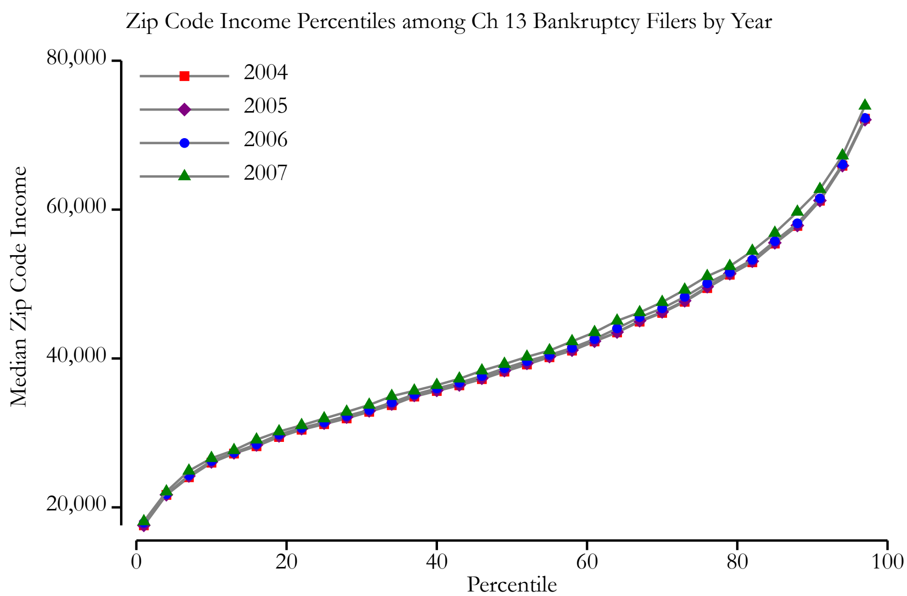
Figure A10: ZIP Code Income




*Notes:* The sample includes all consumer bankruptcy filings included in the PACER sample from January 2004 through December 2007, matched with the ZIP Code median household income measured in the 2000 decennial census. Each dot in the figure represents the average median ZIP Code household income for filers in that week. The vertical line indicates the date when BAPCPA was implemented, October 17, 2005.

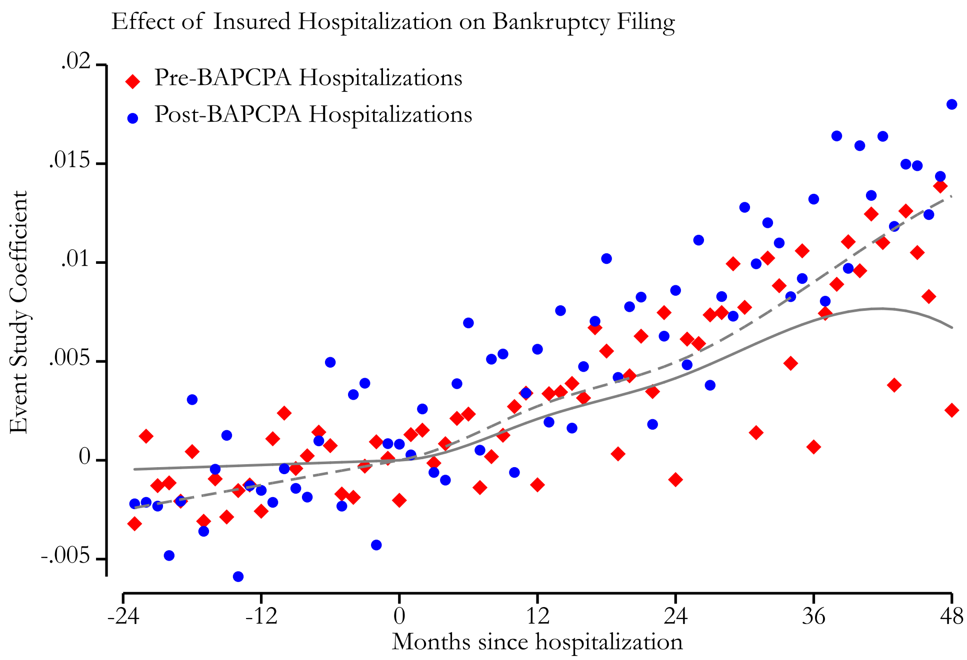
Figure A11: Income Distribution Filers by Chapter



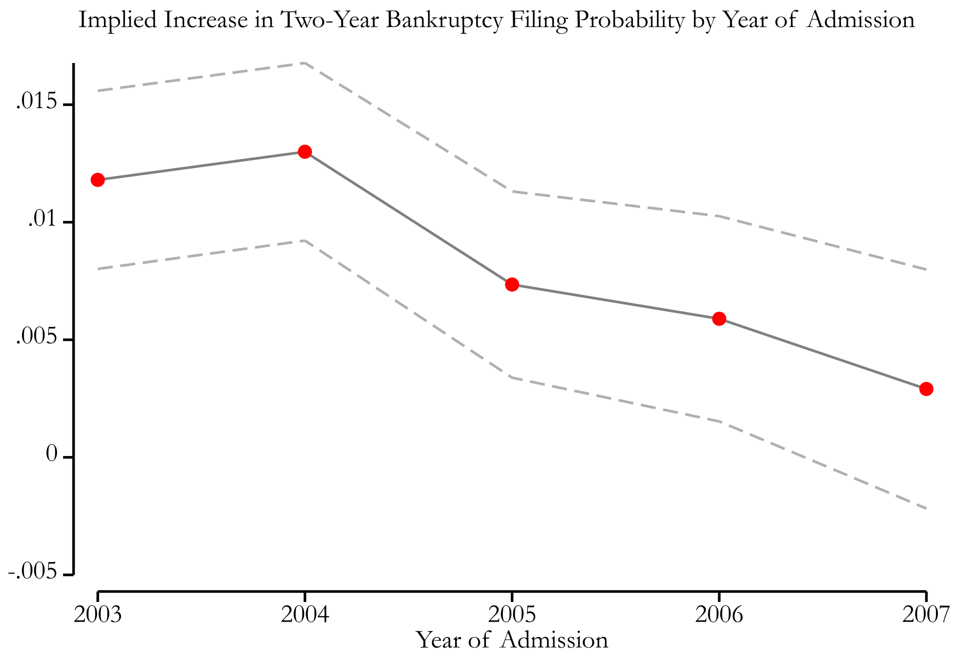

*Notes:* The sample includes all Chapter 7 and 13 consumer bankruptcy filings included in the PACER sample in 2004 and 2006, matched with the ZIP Code median household income measured in the 2000 decennial census. The two distributions plot the percentiles of ZIP Code median household income among filers in 2004 and 2006.

Figure A12: Effect of Hospitalization on Bankruptcy Filing for Insured




*Notes:* The sample is individuals ages 25-64 who are hospitalized with insurance in California, additionally split by the timing of the hospitalization (January 2003 through December 2004 for the pre-BAPCPA sample, October 2005 through December 2007 for the post-BAPCPA sample). The points represent the estimated effects of event time (i.e., the $\mu_r$s from the non-parametric event study in equation 5) and the lines represent the parametric event study in equation 6 with the pre-trends normalized between the two periods for ease of visual comparison.

Figure A13: Effect of Hospitalization on Bankruptcy Filing by Year of Admission




*Notes:* The sample is individuals ages 25-64 who are hospitalized without insurance in California, additionally split by the year of admission to the hospital. The points represent the estimated 24-month implied effects of the hospitalization on the likelihood an individual has filed for bankruptcy. Implied effects are estimated based on the parametric event study in equation 6 with year fixed effects jointly estimated and event study coefficients separately estimated by year of admission.