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Trends in Private Patent Costs and Rents for Publicly-Traded United States Firms

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Abstract

We use detailed data to estimate the private costs and private rents of United States patents for publicly-traded firms. In analyzing costs, we first introduce a novel theoretical model to interpret our estimates. We then combine lawsuit data from Derwent Litalert with non-practicing entity (NPE) lawsuits collected by Patent Freedom, and use an event-study approach to estimate losses suffered by alleged infringers during 1984-2009. To estimate rents, we combine patent data from the USPTO and EPO with financial data from COMPUSTAT, and use market-value regressions to estimate the value of patent rents for publicly-traded US firms during 1979-2002. We find that private costs exceed private rents during 1999-2000 and the trend in costs is sharply higher. Costs also exceed forecasts of rents for 2005-09. A surge in the number of NPE lawsuits contributes to the increase in the gap.

JEL Code: K2, O3
Keywords: event study, market-value regression, patent, litigation, non-practicing entities, research and development

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1. Introduction

Innovation is a major driver of economic growth and is among the most important competitive factors determining profitability and market value. Because benefits from investments in research and development often spill over to non-investing parties (Arrow 1962; Nordhaus 1969), competitive markets may not fully reward innovators for their investments. A patent grants an exclusive right to exploit an invention for a limited time, allowing a patentee to obtain market power, and earn economic "rents," in markets for products and technology. Ideally, expected rents move innovation incentives toward the socially optimal level.

If rents from patents are below private costs generated by patents, however, then it is questionable whether patent rents improve innovation incentives. Bessen and Meurer (2008) argue that the patent system drifted from the principles of property law in the 1980s and 1990s,[1] driving up costs relative to rents. Restricting attention to publicly-traded US firms during 1984-99, they conduct an event study of patent litigation to estimate a lower bound for aggregate private costs of patents, and use market-value regressions to estimate an upper bound for aggregate private rents of patents. They show that by the late 1990s, private costs exceed private rents outside of the chemical and pharmaceutical industries.

We improve the Bessen and Meurer (2008) estimation framework and adapt it to more recent and more comprehensive data. For estimating costs, we first introduce a novel theoretical model of litigation and settlement that helps measure and interpret the effects of litigation filings on firm value. We then extend litigation data to cover 1984-2009, and include lawsuit filings from Derwent Litalert and non-practicing entity (NPE) lawsuits collected by Patent Freedom. Litigating firms are matched to stock market returns from CRSP and additional data in COMPUSTAT. For estimating rents, we match 1969-2002 data on US and European patents to financial data from COMPUSTAT. In contrast to previous studies, our analysis controls for (and estimates rents for) European patents. This produces more precise rent estimates for US patents. Following Bessen and Meurer, we estimate a lower

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[1]Related work discusses how changes to patent law and precedent since the establishment of the Court of Appeals for the Federal Circuit (CAFC) in 1982 led to increased patenting and litigation thereafter (Hall and Ziedonis 2001; Jaffe and Lerner 2004; Hall 2005; Henry and Turner 2006).

bound for aggregate costs and an upper bound for aggregate rents.

Our theoretical litigation model frames expectations for investors purchasing the stocks of publicly-traded firms that face risks of being sued for patent infringement. It specifies patent rents, litigation costs and transfer payments for a patentee and an alleged infringer in three sets of states: no litigation, litigation followed by settlement, and litigation to a decision. It also specifies transition probabilities between states. If settlement occurs, then the patentee and alleged infringer engage in Nash bargaining.

We use the model to show why event studies of the effects of patent litigation filings on firm values tend to underestimate litigation costs. Intuitively, market returns incorporate investors' beliefs that firms may settle litigation. When firms settle via Nash bargaining, they choose royalties to maximize joint patent rents. If the probability of settlement is high, then the expected increase in rents dampens the negative effect of litigation.

We estimate an average per-event loss of $41.4 million, and aggregate losses of $308 billion, for publicly-traded alleged infringers sued for patent infringement during 1986-2009. We construct these estimates using a standard event study to recover a 5-day cumulative abnormal return (CAR) for each event, and a linear regression model relating CARs to firm and event characteristics. Among three alternatives, this model produces the most conservative estimate of average losses.

For patentees, litigation filing has no significant effect on firm value. Our theoretical model then implies that the entire effect of litigation is captured by the effect on alleged infringers. And given that patent infringement suits have very high settlement rates,[2] the aggregate effects on alleged infringers are appropriately interpreted as an upper bound for litigation costs.

Next, consider rents. When we control for European patents, we estimate an average US patent rent of $351,000 for publicly-traded US firms during 1979-2002. When European patents are omitted, then US patents proxy for European patents (and others) and our estimate rises to $517,000. Hence, controlling for European patents lowers the average US

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[2]For example, data from the Federal Judicial Center show that trial frequency in patent infringement suits is about 5.0% during 1996-2000 (the middle years of our data) and declines across time [source: Henry and Turner (2016), Table A1, p. 483].

patent rent estimate by nearly one third.

It is well known that, regardless of whether all foreign patents are included in the estimation, market-value regressions produce upwardly-biased estimates of patent rents. Intuitively, successful firms tend to be both valuable and to yield high innovation quantity and quality from their R&D. We therefore interpret our estimate of $351,000 as an upper bound US patent rent, which is less exaggerated due to our inclusion of European patents. Forecasting annual patent rents over 2003-09 to grow at the same average as 1979-2002 and aggregating estimated rents over 1986-2009, we estimate an upper-bound for total rents of about $342 billion.
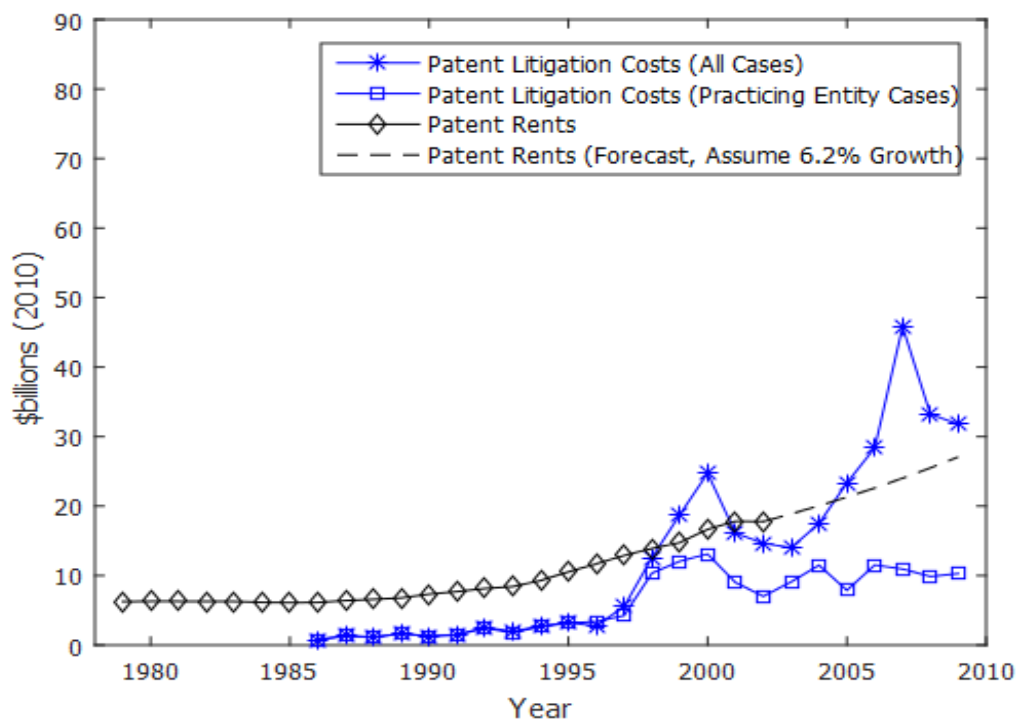
Figure 1: *Estimated Patent Litigation Costs and Rents, Publicly-traded US Firms*

Hence, estimated aggregate rents are slightly higher than estimated aggregate costs. An intertemporal comparison of costs and rents paints a different picture, however. Figure 1 shows that annual growth in costs (16.2%) sharply exceeds growth in rents (6.2%). As a result, estimated costs are higher than estimated rents during 1999-2000 and exceed forecasted rents during 2005-09. NPE cases make the difference. Costs just from practicing entity (PE)

cases never exceed rents.

Consistent with past applications of the event-study methodology (e.g., Bessen and Meurer 2008), our estimated average cost is more than ten times higher than recent estimates of costs of legal counsel (e.g., AIPLA 2011). Our theoretical model specifies both direct and indirect litigation costs, and for some parts of our data we can estimate indirect costs. Bessen and Meurer (2014) estimate the average direct legal cost (cost of legal counsel plus licensing fees) in NPE cases over 2005-11. Using this, we estimate that indirect costs are about 85% of total costs for NPE cases during 2005-09. Interestingly, this varies substantially with firm size. For alleged-infringer firms with less than $1 billion in annual revenue, indirect costs are just 8% of the total. For larger alleged infringers, indirect costs are about 86%.

Within our model, we attribute the high indirect costs to uninsurable business risks of patent litigation. Recent research has shown that firms sued for patent litigation significantly curtail innovative activity (Tucker 2014; Cohen, Gurun and Kominers 2014) and, due to the lower level of innovation, lose demand for their product lines (Tucker 2014). We use a stylized example to identify assumptions that make our cost estimates consistent with the results from these papers. We find that a quite natural set of assumptions—5% yearly discounting, one year of litigation, and constant flow profit that is proportional to demand—generates percentage losses consistent with averages from the event study.

The results from our linear regression model give additional support for the argument that business risk of litigation drives losses. First, we find that percentage losses vary with firm size. Most notably, firms with below $100 million in revenue lose more than 1% per event, while firms with higher revenue lose no more than 0.15% per event. Small firms are more likely to have a small number of product lines, and a single patent lawsuit could destroy a small firm. For larger firms, patent disputes are more of a routine part of doing business.

Next, we show that when an alleged infringer has not been involved in patent litigation in the previous two years, it loses 0.21% more value when sued. For all but the smallest alleged infringers, costs are more than doubled in such "surprise" lawsuits. The magnitude of the effect suggests that business risk drives costs. While unprepared firms may need to scramble

to put together a viable defense, it seems unlikely that this would increase direct litigation costs by enough to double overall costs. On the other hand, a firm's lack of preparation could cause it to make more mistakes in adjusting its innovation and product-release strategies, à la Tucker (2014) and Cohen et al. (2014), to a patent-litigation environment. In addition, for firms frequently involved in litigation, a given lawsuit is more likely to involve a repeat rival. This experience should make it easier to avoid major strategic business mistakes or to quickly settle the dispute.

Some researchers have argued that suits involving NPEs tend to introduce particularly costly hold-up problems, even when patents may be weak (Shapiro 2010).[3] We find that alleged infringers that face an NPE lose just as much as firms who do not. Because NPE cases involve more defendants, however, our results suggest that losses per suit are worse than average.

Like Bessen and Meurer, we do not estimate the *social* costs and rents of the patent system, just the private costs and rents. If overall private costs exceed private rents, then it is unlikely that the patent system is providing positive innovation incentives overall. Furthermore, we consider only publicly-traded firms, so our conclusions regarding incentives apply only to this group. While publicly-traded firms are responsible for most research and development spending in the US,[4] some commentators argue that small private firms are a more important source of innovation. Bessen and Meurer (2008) examine this issue more fully and conclude otherwise, and new evidence suggests that much of the patent litigation in recent years from NPEs is, in fact, targeted at small companies (Bessen and Meurer 2014). Hence, they bear significant litigation costs as well.

2. Literature Review

The event study method dates to Fama, Fisher, Jensen and Roll (1969). The first paper to use an event study to estimate the costs of litigation is Cutler and Summers (1988),

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[3]See particularly the discussion of the infringement case between NTP, Inc. and Research in Motion (Shapiro 2010, p. 281).

[4]For example, during 1999 US public firms spent $150 billion on R&D, while total company spending on US industrial R&D was $160 billion. See Bessen and Meurer (2008, p. 142) for a detailed discussion of how to interpret these numbers, which are based on National Science Foundation R&D statistics.

who study litigation between Texaco and Pennzoil over mergers. For patent litigation more specifically, Bhagat, Brickley and Coles (1994), Lerner (1995) and Bhagat, Bizjak and Coles (1998) study lawsuit announcements in the *Wall Street Journal* or similar periodicals and estimate the combined loss of plaintiffs and defendants to be between 2-3% of value.[5] Bessen and Meurer (2012) find similar results for a slightly larger sample of cases. The average CAR estimated in these studies is significantly higher than our estimates. In all of these instances, however, the number of observations is very small and highly selected.

Bessen and Meurer (2008; 2012) also consider a large cross section of case filings. Using an untrimmed sample of lawsuits from Derwent Litalert for 1984-99 and a 25-day window, Bessen and Meurer estimate an average loss of 0.50% for alleged infringers.[6] Among the sample firms, this implies a mean loss of $75.9 million and a median loss of $6.5 million.[7] In comparison to these past studies, our approach produces more conservative average (per event) cost estimates. Most importantly, we use a 5-day window and trim out events where there is an overlap with another event involving the same firm. The shorter window leaves out some costs, while the trimming eliminates observations where event overlaps could lead to double-counting of losses. We also include events from 2000-09, which have lower average CARs as well. Bessen, Meurer and Ford (2011) analyze an untrimmed sample of Patent Freedom data and estimate an average loss of about 0.37% for NPE lawsuits over 1990-2010. This implies aggregate losses for defendants in NPE cases of about $579 billion.[8]

The literature on patent rents starts with Griliches (1981), who finds a significant relationship between firm market value and intangible capital, as proxied by past R&D expenditures and patent counts. Subsequent studies adapt the Griliches (1981) framework to show that numbers of patents (Deng, Lev and Narin 1999; Bosworth and Rogers 2001), the ratio of patents to R&D (Hall, Jaffe and Trajtenberg 2005), patent citations (Shane and Klock 1997;

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[5]See Lunney (2004) and Haslem (2005) for other studies of the effects of lawsuit announcements. Henry (2013) uses data from published decisions to show that firms whose patents are invalidated lose about 1% of value. Panattoni (2011) shows that branded pharmaceutical firms that lose Paragraph IV cases (thereby permitting generic entry) suffer sizable losses.

[6]The first to use Derwent Litalert data, Lanjouw and Schankerman (2001) analyze a wide number of characteristics of litigated patents.

[7]This is in 2010 dollars and is reported in Table 3 of Bessen, Ford and Meurer (2011).

[8]See our online appendix (http://people.terry.uga.edu/jlturner/BNTWAppendixNovember2017.pdf) for estimates of average and aggregate costs using samples comparable to those used by Bessen and Meurer (2008; 2012) and Bessen, Meurer and Ford (2011).

Deng, Lev and Narin 1999; Hall, Jaffe and Trajtenberg 2005; Chin et al. 2006; Lee 2008; Chen and Chang 2010; Chen and Shih 2011) and technological diversity (Miller 2006) are associated with higher firm value.[9] However, these papers do not formally model how patent rents affect firm value. With such a model, Bessen (2009b) estimates an upper bound patent rent of $536,000 (2010 dollars). Our paper follows Bessen (2009b), but also controls for European patent stocks and estimates a lower average rent.

The renewal method (Pakes and Schankerman 1984; Pakes 1986) is the main alternative model for estimating patent rents. A firm that pays renewal fees to maintain a patent holds a more valuable patent than a firm that does not pay. Exploiting variability in the size of fees and the timing of renewal decisions, this method imputes a distribution of the discounted value of rents to estimate mean rents. Bessen (2008) estimates a mean patent rent of about $78,000 for US patents issued in 1991. In related work, Putnam (1995) also uses observed patentee behavior, in the form of decisions to file patents in foreign countries, to estimate patent value. Serrano (2005) similarly uses decision to sell patents to impute value.

Unfortunately, the renewal method requires parametric assumptions (usually log-normality) to identify mean patent rents. But there is considerable evidence that upper tails of patent value distributions are thicker than tails for log-normal distributions. Harhoff, Scherer and Vopel (2003) argue the distribution of rents is Pareto, and with parameter values such that mean rents are infinite. Fortunately, Bessen (2009b) shows that estimates of patent rents using the market-value method do not indicate infinite rents.[10] Renewal studies also have a severe right-censoring problem; one must observe the renewal history of a cohort of patents to estimate their value. This means that there is a long lag between when the value is learned by innovators and when it is learned by researchers.

Of course, the real question is whether patents generate more costs than rents. In work subsequent to the Bessen and Meurer (2008) comparisons, Hall and MacGarvie (2010) an-

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[9]Sandner and Block (2011) also show that the market value of publicly traded companies, as measured by Tobin's Q, is positively related to the registration of trademarks.

[10]More formally, if the distribution of value is Pareto, then for particular values of the parameter, the mean is undefined. Thus, estimates of the mean using repeated sampling off of this distribution would not converge. Scherer (1998) presents evidence suggesting that while the variance is undefined, the mean probably is defined. Bessen (2009b) shows that for US firms during 1979-97, increasing the patent portfolio size does improve the precision of estimates.

alyze patenting by firms in the ICT sector. They find that the market evaluates software patents negatively on aggregate, but firms that patent have slightly higher market values than those that do not. Hence, private incentives to patent are positive but social surplus is negative.

3. A Model of Patent Litigation

Consider a patent-owning firm that has a rival. In the course of competition between the patentee and the rival, investors learn information that permits the formation of expectations of future profits that account for numerous states of the world. These states include no patent litigation, filing of patent litigation followed by settlement, and full-blown patent litigation with possible assignment of royalty damages.

We introduce a stylized investor-centric model, illustrated in Figure 2, to organize the various states and to allow us to characterize how expected payoffs change when litigation is filed. The pre-litigation information set available to investors is held at point A at the top of the tree, and it includes the payoffs down all branches of the tree, and the transition probabilities. At point B, investors possess all of the information that they had at point A, and also know that litigation has been filed.

At point A, investors do not know the full nature of the relationship between the firms, and expect one of two states to emerge. In the "imitation" state, which occurs with probability λ, the rival imitates the patentee's technology and harms its payoff (conditional on no litigation). The patentee then finds it optimal to litigate. The "no imitation" state occurs with probability $1 - \lambda$, and does not induce litigation. Hence, investors have correct expectations over the probability of litigation.

We will also refer to the rival (interchangeably) as the alleged infringer. Prior to any resolution of litigation, let the probability that the patentee wins an infringement suit be π.[11] We will interpret this as an index of the patent's overall strength. Let any per-unit royalty paid by the alleged infringer to the patentee be r, and let any lump sum payment be

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[11]For simplicity, we ignore any uncertainty about infringement, and think of uncertain court outcomes as reflecting uncertainty over patent validity.
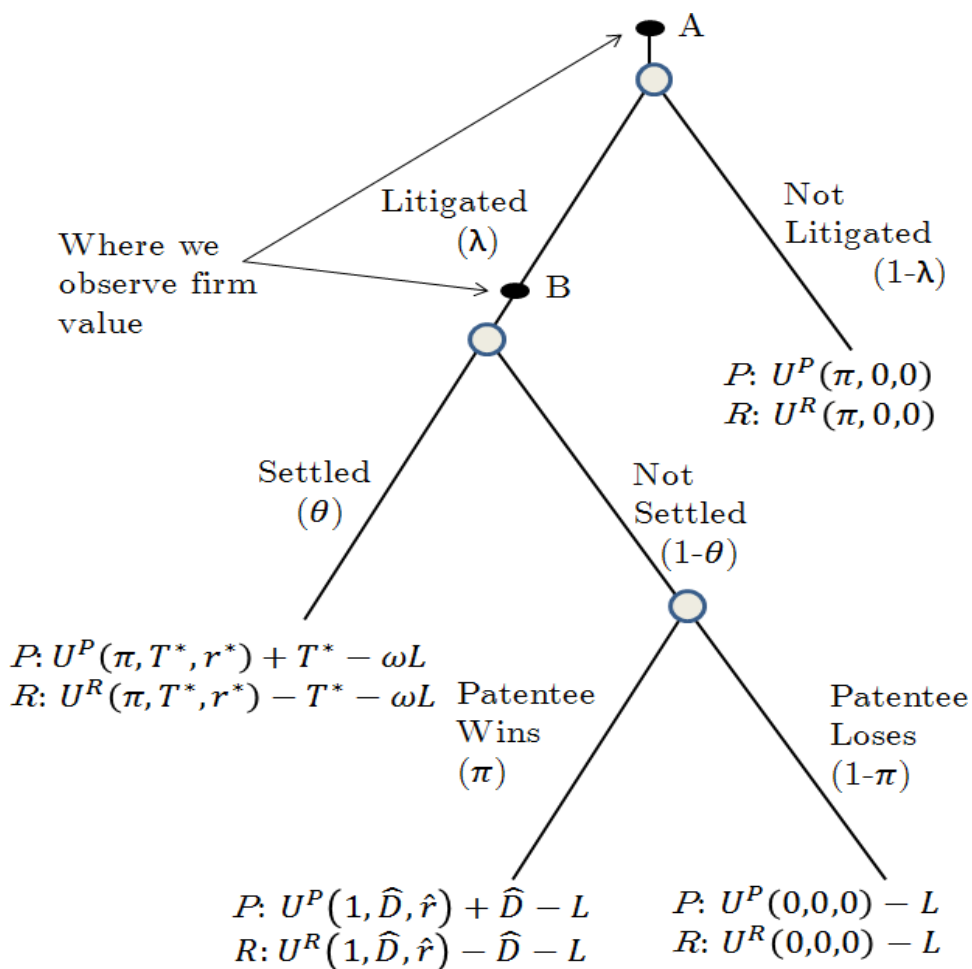
Figure 2: *Litigation Game Tree*

T. Let all payoffs that do not depend on the patent be P_0 for the patentee and R_0 for the rival, and for simplicity normalize these payoffs to zero.

By virtue of having the patent, the patentee earns profits that are no lower than it otherwise would earn. Call these extra profits *rents* and denote them as $U^P(\pi, T, r)$. Intuitively, we will think of rents as depending upon the overall strength of the patent π, and on royalties T and r that obtain either due to settlement or due to a court decision. Note also that we define utility net of any transfer payments from the defendant, so it is best to think of them as all rents other than pure transfers from the defendant.

For the defendant, denote its rents as $U^R(\pi, T, r)$. The second and third arguments reflect the fact that other rivals may respond to the levels of T and r. Note that r captures all aspects of any bargain other than T, including whether there is a settlement.

If litigation does not occur, then the expected payoffs are:

$$P : V^P_{NoFile} = U^P(\pi, 0, 0)$$
$$R : V^R_{NoFile} = U^R(\pi, 0, 0)$$

These *baseline* reflect states of the world where the patentee does not find it optimal to sue.[12] If litigation is filed, then the patentee and the alleged infringer bargain over a settlement in the shadow of possible litigation. For simplicity we restrict attention to one round of 50/50 Nash bargaining followed (possibly) by one simple litigation subgame. If the firms do not settle, then each firm pays direct litigation cost L_D and indirect litigation cost L_I for full-blown litigation. Let the total litigation cost be $L \equiv L_D + L_I$.

In pursuing litigation, the patentee and alleged infringer must pay direct costs to prepare for settlement negotiation and possible courtroom activity. These include costs of legal counsel and of discovery. Indirect costs include business risks of patent litigation. While litigation is ongoing, the patentee and alleged infringer may find it optimal to adjust investments in innovation and new product releases. For example, the alleged infringer may prefer not to release new products prior to completing litigation, out of fear of being sued for willful infringement. L_I also includes a firm's opportunity cost of committing personnel to assist with litigation, instead of doing other things.

With probability π the patentee wins, and the court assigns payment $\widehat{D}$ and running royalty rate $\widehat{r}$. The payoffs are:

$$P : U^P(1, \widehat{D}, \widehat{r}) + \widehat{D} - L$$
$$R : U^R(1, \widehat{D}, \widehat{r}) - \widehat{D} - L$$

With probability $1 - \pi$ the patentee fails to win damages, owns a worthless patent and receives no rents. The payoffs are:

$$P : U^P(0, 0, 0) - L$$
$$R : U^R(0, 0, 0) - L$$

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[12]In a more general, dynamic, model, baseline rents would reflect states of the world where the patentee does not *yet* find it optimal to sue, and future states could include cases where infringement occurs and litigation follows.

Hence, the expected payoffs conditional on litigation are:

$$P : V_{Lit}^P = \pi \left[U^P(1, \widehat{D}, \widehat{r}) + \widehat{D} \right] + (1 - \pi)U^P(0, 0, 0) - L$$
$$R : V_{Lit}^R = \pi \left[U^R(1, \widehat{D}, \widehat{r}) - \widehat{D} \right] + (1 - \pi)U^R(0, 0, 0) - L$$

Prior to litigation, the firms bargain over terms of a settlement. With probability θ, litigation costs under settlement equal ωL, with $\omega < 1$, and settlement is efficient. With probability $1 - \theta$, litigation costs under settlement are so large that settlement is inefficient and litigation ensues. Hence, the game played by the firms leads to the exact probability of settlement that investors expect.

If the terms of the settlement are such that the alleged infringer agrees to pay royalties according to r and fixed payment T, then the payoffs are

$$P : U^P(\pi, T, r) + T - \omega L$$
$$R : U^R(\pi, T, r) - T - \omega L$$

In an equilibrium Nash bargain, the firms choose r and T to maximize their *joint* rents. Moreover, each firm earns its expected payoff under litigation (including expected damages) plus one half of the extra surplus earned from settlement versus litigation.

By reaching a bargain with terms r and T, the parties avoid litigation and realize additional bargaining surplus

$$\Delta S = \ U^P(\pi, T, r) - \left[\pi U^P(1, D, r) + (1 - \pi)U^P(0, 0, 0) \right]$$
$$+ \left\{ U^R(0, T, r) - \left[\pi U^R(1, D, r) + (1 - \pi)U^R(0, 0, 0) \right] \right\} + 2(1 - \omega)L.$$

The terms on the first line give the expected additional total rents accruing to the patentee. The terms in brackets on the second line give the expected additional total rents accruing to the alleged infringer. The third term is savings on litigation costs. Under Nash bargaining, the patentee and alleged infringer maximize and then split this surplus. In equilibrium, the royalty equals r^*, the fixed payment equals T^*.

Conditional on an efficient bargain, this leads to the following payoffs:

$$
\begin{aligned}
V_{Settle}^P = \quad & \pi\widehat{D} - \omega L + \tfrac{1}{2}\Big\{ U^P(\pi, T^*, r^*) + \pi U^P(1, \widehat{D}, \widehat{r}) \\
& + (1-\pi)U^P(0,0,0) + U^R(0, T^*, r^*) - \Big[\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi)U^R(0,0,0)\Big]\Big\} \\
V_{Settle}^R = \quad & -\pi\widehat{D} - \omega L + \tfrac{1}{2}\Big\{ U^R(0, T, r) + \pi U^R(1, \widehat{D}, \widehat{r}) \\
& + (1-\pi)U^R(0,0,0) + U^P(\pi, T^*, r^*) - \Big[\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi)U^P(0,0,0)\Big]\Big\}.
\end{aligned}
$$

Conditional on the filing of a lawsuit, expected valuations are:

$$
\begin{aligned}
P &: V_{File}^P = \theta V_{Settle}^P + (1-\theta)V_{Lit}^P \\
R &: V_{File}^R = \theta V_{Settle}^R + (1-\theta)V_{Lit}^P.
\end{aligned}
$$

Recall Figure 2. At point A, uncertainty about the nature of competition has not yet been resolved. Hence, ex ante expected payoffs are:

$$
\begin{aligned}
P &: V_{ExAnte}^P = \lambda V_{File}^P + (1-\lambda)V_{NoFile}^P \\
R &: V_{ExAnte}^R = \lambda V_{File}^R + (1-\lambda)V_{NoFile}^R.
\end{aligned}
$$

When the patentee decides to file litigation, the expected payoff for firm i changes by $\Delta V^i = V_{File}^i - V_{ExAnte}^P$. We have:

$$
\begin{aligned}
\Delta V^P = \quad & (1-\lambda)\Big\{ \pi\widehat{D} + \tfrac{2-\theta}{2}\Big[\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi)U^P(0,0,0)\Big] - U^P(\pi, 0, 0) \\
& + \tfrac{\theta}{2}\Big[U^R(0, T^*, r^*) + U^R(\pi, T^*, r^*) - (\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi)U^R(0,0,0))\Big] \\
& - L(\theta\omega + 1 - \theta)\Big\} \\
\Delta V^R = \quad & (1-\lambda)\Big\{ -\pi\widehat{D} + \tfrac{2-\theta}{2}\Big[\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi)U^R(0,0,0)\Big] - U^R(\pi, 0, 0) \\
& + \tfrac{\theta}{2}\Big[U^R(0, T^*, r^*) + U^P(\pi, T^*, r^*) - (\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi)U^P(0,0,0))\Big] \\
& - L(\theta\omega + 1 - \theta)\Big\}
\end{aligned}
$$

(1)

These expressions highlight what the event study picks up. First, expected damages are $\pi\widehat{D}$ for the patentee and $-\pi\widehat{D}$ for the defendant. Although litigation is carried through to a decision with just probability $1-\theta$, the patentee earns expected damages in all scenarios—as a threat point under settlement, or as an expected damage payment in litigation. This is the only part of payoffs where the magnitude of the effect of litigation is unambiguously different for the patentee and defendant.

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Second, part of firm value changes because expected rents change. For the patentee and defendant, these are, respectively:

$$P : \tfrac{2-\theta}{2} \left[\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi) U^P(0,0,0) \right] + \tfrac{\theta}{2} U^P(\pi, T^*, r^*) - U^P(\pi, 0, 0)$$
$$R : \tfrac{2-\theta}{2} \left[\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi) U^R(0,0,0) \right] + \tfrac{\theta}{2} U^R(0, T^*, r^*) - U^R(\pi, 0, 0).$$

Note that the nature of Nash bargaining payoffs dampens the effect that settlement has on rents. If $\theta = 0$, so that litigation is certain, then these payoffs are just expected rents under litigation minus status-quo rents under "no litigation." If $\theta = 1$, on the other hand, then settlement is certain and the patentee rents created through the bargain are shared with the defendant, so the patentee gets just one-half of those rents.

Third, part of each firm's value changes because the firm expects to earn some share of the change in rents from the *other firm*. They are:

$$P : \tfrac{\theta}{2} \left\{ U^R(0, T^*, r^*) - \left[\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi) U^R(0,0,0) \right] \right\}$$
$$R : \tfrac{\theta}{2} \left\{ U^P(\pi, T^*, r^*) - \left[\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi) U^P(0,0,0) \right] \right\}.$$

Generally, we expect this change to be second-order.

Finally, each party expects to spend resources. These are:

$$P : -L(\theta\omega + 1 - \theta)$$
$$R : -L(\theta\omega + 1 - \theta).$$

For both firms, value changes due to expected litigation costs. This is high either if settlement is unlikely (θ is low) or if settlement is costly (ω is high).

Because of the asymmetry due to expected damages, we expect $\Delta V^P > \Delta V^R$. But it is possible for ΔV^P to be negative, zero or positive. The reason is the model treats the filing of litigation as revealing previously unknown information about the true state of competition. For litigation to be privately optimal in the imitation state, the patentee's expected payoff when filing litigation V_{File}^P must exceed its hypothetical payoff under no litigation for the imitation states. For no litigation to be privately optimal in the no imitation state, the patentee's payoff under no litigation $U^P(\pi, 0, 0)$ must exceed its hypothetical payoff under litigation in the no imitation state. The relationship between $U^P(\pi, 0, 0)$ and V_{File}^P is

ambiguous, and ΔV^P will tend to be higher when imitation more dramatically increases the payoff to litigation.[13]

Importantly, if $\Delta V^P = 0$, then the entire effect of litigation shows up in ΔV^R. Imposing $\Delta V^P = 0$ onto the first expression in (1), then substituting into the second expression, we identify this effect.

Proposition 1. *Suppose $\Delta V^P = 0$. Then*

$$
\begin{aligned}
\Delta V^R = \ & (1-\lambda)\left\{-2L(\theta\omega + 1 - \theta)\right. \\
& + \theta U^P(\pi, T^*, r^*) + (1-\theta)\left[\pi U^P(1, \widehat{D}, \widehat{r}) + (1-\pi)U^P(0,0,0)\right] - U^P(\pi,0,0) \\
& \left. + \theta U^R(0, T^*, r^*) + (1-\theta)\left[\pi U^R(1, \widehat{D}, \widehat{r}) + (1-\pi)U^R(0,0,0)\right] - U^R(\pi,0,0)\right\}
\end{aligned}
$$
(2)

The first term inside brackets is the total litigation costs faced by *both* firms. The terms on the second line give the difference in expected rents for the patentee from "no litigation" versus the expected rents at the point litigation is initiated. The terms on the third line give the analogous difference for the expected rents for the alleged infringer.

Proposition 2. *Let the set of joint rents that the patentee and alleged infringer bargain over, after litigation is filed, include the joint rents under the baseline. If baseline joint rents are not strictly maximal and if $\Delta V^P = 0$, then for sufficiently high θ, we have $\Delta V^R > -(1-\lambda)2L(\theta\omega + 1 - \theta)$. That is, ΔV^R understates expected litigation costs.*

Proof of Proposition 2. Because r and T maximize joint rents, these rents must exceed the joint status quo rents, i.e., $U^P(\pi, T^*, r^*) + U^R(\pi, T^*, r^*) - [U^P(\pi,0,0) + U^R(\pi,0,0)] > 0$. For $\theta = 1$, we have

$$
\begin{aligned}
\Delta V^R \ = \ & (1-\lambda)\left\{-2L\omega \right. \\
& \left. + U^P(\pi, T^*, r^*) + U^R(\pi, T^*, r^*) - [U^P(\pi,0,0) + U^R(\pi,0,0)]\right\} \\
> \ & (1-\lambda)(-2L\omega).
\end{aligned}
$$

By continuity, $\Delta V^R > -(1-\lambda)2L(\theta\omega + 1 - \theta)$, for smaller values of θ. **QED.**

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[13]To see this more clearly, let the payoff to suing in the no imitation state be $V^P_{File,NoIm}$. We know that $U^P(\pi,0,0) > V^P_{File,NoIm}$. Hence, if $V^P_{File} > V^P_{File,NoIm}$, then V^P_{File} may (or may not) exceed $U^P(\pi,0,0)$.

Intuitively, when litigation is filed, the values of the firms incorporate expected future litigation costs and the expected change in future rents. The change in rents must be positive as long as settlement is sufficiently likely, because settlement maximizes joint rents. For the limiting case where baseline rents happen to be jointly optimal, settlement merely reinforces baseline rents. Then, ΔV^R exactly equals $(1 - \lambda)$ times litigation costs when $\theta = 1$. It may overstate $(1 - \lambda)$ times costs when $\theta < 1$, because expected rents under litigation could be below the (jointly optimal) baseline. Then, ΔV^R will understate litigation costs provided the ex ante probability of litigation λ is sufficiently high.

Thus far, we have treated all litigation costs as being symmetric. However, it seems more realistic that (in the absence of a counter-suit) the patentee and alleged infringer will face asymmetric indirect costs. In particular, if the threat of willfulness claims drive the decisions that determine indirect costs, then the alleged infringer will face far higher indirect costs than the patentee. Suppose indirect costs remain L_I for the alleged infringer but are approximately zero for the patentee. Then we have, for $\theta = 1$,

$$\begin{aligned}
\Delta V^R = \quad & (1 - \lambda)\left\{-(2L_D + L_I)\omega \right. \\
& \left. + U^P(\pi, T^*, r^*) + U^R(\pi, T^*, r^*) - [U^P(\pi, 0, 0) + U^R(\pi, 0, 0)]\right\}
\end{aligned}$$

For this case, the event study yields an upper-bound estimate for the combined direct litigation costs of both firms, plus the alleged infringer's indirect litigation costs.

4. Estimation of Costs

Assuming the efficient market hypothesis holds, the event-study method estimates the average effect of litigation by measuring the market-value reaction to litigation filing. Following Salinger (1992), we specify the following market model:

$$\rho_{it} = a + b\rho_{it}^m + \epsilon_{it},$$

where ρ_{it} is the return to stock i on day t, ρ_{it}^m is the compounded return on the CRSP value-weighted market index and ϵ_{it} is a mean-zero error term. The market portfolio is included as a regressor to filter out that part of the return that is not due to the litigation filing.

Now, consider an event that occurs on day T. The following model permits a regression of "abnormal" returns on that day:

$$\rho_{it} = a + b\rho_{it}^m + \psi I_{it} + \epsilon_{it}, \tag{3}$$

where $I_{it} = 1$ if $t = T$ and 0 otherwise. We estimate this model for event i, by ordinary least squares regression, using 200 trading days and ending two trading days prior to the event.[14] We consider a 5-day event window (from -1 to +3). The cumulative abnormal return (CAR) is the sum of all of the daily abnormal returns during the event window. This exercise is repeated for all events.

4.1. The Data

Our litigation data combine case filings from Derwent Litalert with NPE lawsuits captured by Patent Freedom. Derwent Litalert, which includes litigated patents for 1975-present, is available through WESTLAW.[15] Federal courts are required to report all lawsuits filed that involve patents to the US Patent and Trademark Office, and Derwent's data are based on these filings. It is appropriate to think of Derwent data as a random sample with a rate of sampling that varies across time.

Patent Freedom defines an NPE as a company that does not "practice its invention in products or service, or otherwise derive a substantial portion of their revenues from the sale of products and services in the marketplace. Instead, NPEs seek to derive the majority of their income from the enforcement of patent rights." The Patent Freedom data reflect a search of court records to identify US patent lawsuits involving NPEs. In contrast to Derwent data, these records do not include patent numbers but do classify each case as an infringement suit or declaratory judgment. While the data extend back to 1979, coverage is highly incomplete prior to 1990.

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[14]If we have between 150 and 200 trading days' worth of data, then we estimate the model with the data and include the event. If we have fewer than 150 trading days' worth of data, then we drop the event.

[15]Because coverage is lower and is very inconsistent during 1975-83 (Lanjouw and Schankerman 2001), we restrict attention to 1984 and after.

4.2. Lawsuits

Our data include all Derwent cases from 1984-2009, as well as 1990-2009 data from Patent Freedom. The main advantages of the Derwent and Patent Freedom data are the size of their cross sections. The Derwent data for 1984-2009 include 35,301 cases and the Patent Freedom data include 3,249 cases. As we discuss below, there is significant overlap between these sets.

We rely on publicly-traded firms for our main analysis. Hence, we match the parties in these lawsuits to public firms and their financial information from COMPUSTAT and CRSP.[16] For identifying publicly-traded firms from each Derwent case, we use the matching algorithm of Bessen (2009a). The set of public firms from the Patent Freedom data is the same as that used by Bessen, Meurer and Ford (2011), except it omits cases from 2010. We match Patent Freedom cases to Derwent cases using filing dates and docket numbers.

Upon completing these matches, we identify 9,478 cases from Derwent and 1,414 cases from Patent Freedom that involve at least one public firm match. Hence, about 27% of all Derwent cases and about 44% of all Patent Freedom cases include at least one matched public firm. Of the 1,414 Patent Freedom cases, 954 match to a Derwent case (about 67%), while 460 do not. Hence, the total number of cases is 9,938. These matches yield a total number of 13,526 firm-lawsuit pairs.

Figure 3(a) shows total lawsuit filings, Derwent lawsuits and total public-firm Derwent lawsuits.[17] For most of 1984-2009, Derwent filings cover 50-70% of all patent lawsuits. Public-firm-match lawsuits are 15-20% of the total. The relatively low frequency of public-firm matches is consistent with the well-documented higher litigation hazard rate among smaller and non-public firms.[18] Figure 3(b) shows public-firm Derwent lawsuit filings for both practicing entity (PE) cases and NPE cases. This Figure does not include Patent

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[16]The Center for Research in Security Prices (CRSP) data, collected and maintained by the Booth School of Business at the University of Chicago, form the most comprehensive collection of security prices, returns and trading volume for the New York Stock Exchange, American Exchange and NASDAQ. These data include uninterrupted time series dating to 1925. CRSP also includes the CRSP/COMPUSTAT Merged Database, which dramatically simplifies matching COMPUSTAT financial data to CRSP security data.

[17]Total case filings are presented by fiscal year, while the other filings are presented by calendar year. The total case filings data come from the Administrative Office of the US Courts: http://www.uscourts.gov/judicialfactsfigures/.

[18]Belenzon, Berkovitz and Bolton (2009) estimate that public firms in the most financially-developed countries obtain just 60% of patents. Hence, public-firm coverage should be well below 100% of Derwent's coverage. Moreover, in their analysis of Derwent data for 1978-99 and a matched set of random patents,

(a) Overall (Unadjusted)	(b) By Patentee Type (Adjusted)
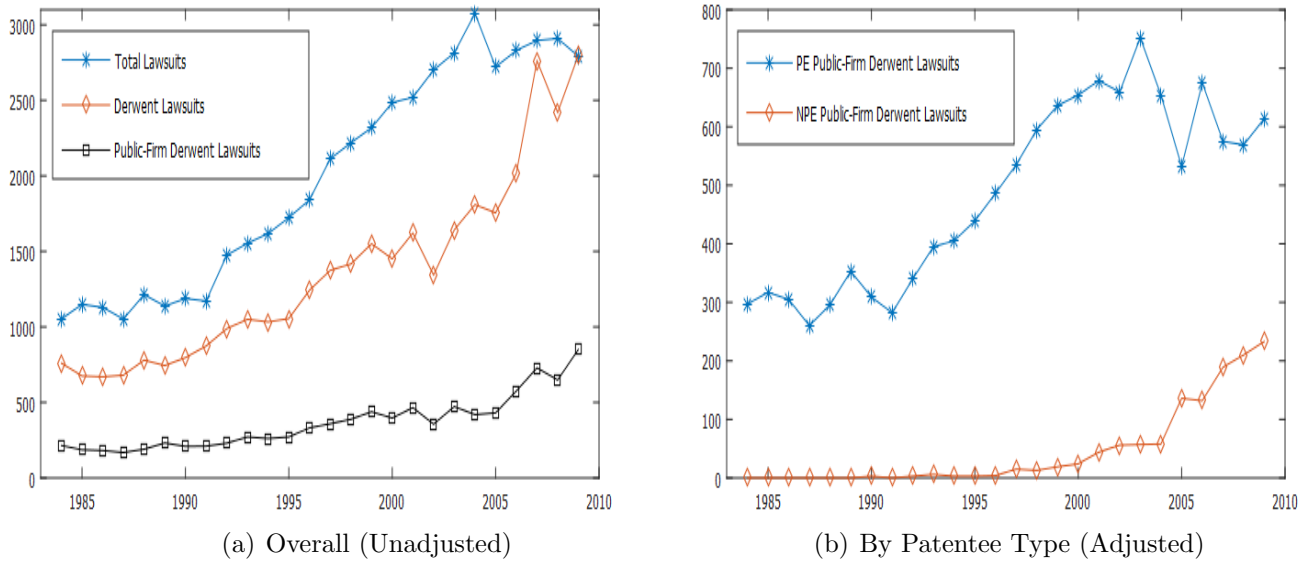
Figure 3: Patent Lawsuits 1984-2009

Freedom lawsuits that do not match to a Derwent lawsuit. We adjust the numbers of cases to account for Derwent's coverage, using 2009 as base year. Strikingly, the number of PE cases peaks in 2003, and is lower in 2009 than in 1999. The number of NPE cases rises sharply after 1999 and explodes after 2004.

In our estimation of aggregate private costs, we do not adjust for coverage as in Figure 3(b) but we do include Patent Freedom lawsuits that do not match to Derwent lawsuits. Hence, our cost estimates miss many litigation events overall, but miss relatively more PE cases. These issues do not hinder our main goal to estimate a lower bound for the aggregate private costs of patent litigation. However, they should be kept in mind when interpreting the share of private costs stemming from NPE cases.

Derwent records include the inventor(s), assignee(s), plaintiff(s), and defendant(s). Comparing the names of the inventors and assignees to the names of the plaintiffs and defendants, we classify cases as infringement suits or declaratory judgments. When the inventor and/or assignee named is the same as the plaintiff, we classify the suit as an infringement suit.

Lanjouw and Schankerman (2004) find that litigation is more than four times more likely for non-public firms.

When the inventor and/or assignee named is the same as the defendant, we classify the suit as a declaratory judgment. When there is no name match—which occurs if the patent is sold between its issue and the lawsuit—we cannot identify the case type.

Table 1 shows descriptive statistics for Derwent cases, while Table 2 describes the Patent Freedom cases. Among all 35,301 Derwent cases, we identify the case type about 66% of the time. Among the 9,478 Derwent cases with at least one public firm, we identify the case type about 73% of the time. The share of infringement suits is higher for the public sample, at about 86% versus 83%. The reason for both discrepancies is that Patent Freedom cases match to only public-firm cases. The Patent Freedom data always identify the case type, and nearly all of these cases are infringement suits.

We also classify the number of defendants in the cases. In records where just one defendant is listed, we classify the case as having a single defendant. In cases with multiple listed defendants, however, Derwent records often describe an ambiguous number of defendants, such as "John Does 1-X."[19] Hence, we classify all cases as having either "one" or "two or more" defendants. Among the full sample of Derwent cases in Table 1, about 79% have a single defendant. The shares are different for the public sample, with 70% single-defendant cases. NPE cases tend to include more defendants, as illustrated in Table 2.

Most Derwent records list a "main" patent, which is classified as either "Design," "Plant," "Reissue" or "Utility." The share of utility patents among public Derwent cases, about 93%, is higher than for the full Derwent data, 90%. Patent Freedom records do not include patent numbers, so we classify patents in NPE cases only for Patent Freedom cases that match to Derwent. These cases overwhelmingly include utility patents (97%).

Following the NBER patent data,[20] we classify patents into one of six technology categories: 1. Chemical, 2. Computers and Communications, 3. Drugs and Medical, 4. Electrical and Electronic, 5. Mechanical, 6. Other. Table 1 shows that the public Derwent cases oversample from the Computers/Communications and Drugs/Medical categories, and undersample from the Mechanical and Other categories. In Table 2, we see that NPE cases

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[19]In addition, Derwent frequently records defendants in the style "XYZ Corp. et al." This problem is particularly bad in later years.

[20]http://www.nber.org/patents/.

overwhelmingly include patents from the Computers/Communications category (77%).[21]

4.3. Events

Each lawsuit includes at least two firms, each of which is affected by the filing of the suit. An event is a firm-lawsuit pair. Consistent with this, we adopt the term "litigation event" when distinguishing events among the characteristics of all lawsuits, and adopt the term "event party" when distinguishing events among the characteristics of all firms involved in those lawsuits. The total number of public firm events is 13,526.

The role an event party plays in a lawsuit is important, because the filing of a suit affects a party differently, depending upon whether it expects to pay or receive damages. We classify event parties as either alleged infringers or patentees. For event parties in infringement suits, we classify plaintiffs as patentees and defendants as alleged infringers. For event parties in declaratory judgments, we classify plaintiffs as alleged infringers and defendants as patentees. If we cannot tell if the suit is for infringement, then we classify defendants as alleged infringers and plaintiffs as patentees. Alleged infringers comprise 8,607 of the event parties (about 64%) while patentees comprise 4,919 of the event parties (about 36%).

Ideally, we would know the precise window during which lawsuits affect firm value. Because most lawsuits are not announced in major publications, however, complete information about the nature of complaints may not be disseminated widely to investors all at once. For these reasons, we run the event study for a 5-day window.[22]

4.4. Trimming the Sample and Event-Party Descriptive Statistics

To identify the CAR of an event, its event window must not overlap with another event window. To see why, suppose a firm gets sued on consecutive days and that the two litigation events are statistically independent, with an average CAR of -0.5%. Then because the event

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[21]One possible explanation is that this is because Computers & Communications patents cover technologies with more complementary components. When this is the case, the incidence of inadvertent infringement is higher, so firms that do not produce are able to acquire more licensing fees and/or damages (Turner 2018).

[22]See our appendix, which reports results using a 25-day window and explains why the 5-day window makes more sense for our data.

windows overlap almost perfectly, the event-study regressions will double count some losses for both events and estimate the average (negative) effect of litigation to be higher than 0.5%. On the other hand, if the events are not statistically independent, this effect could go in the opposite direction. This problem is more severe, the longer is the event window.

Unfortunately, a significant share of our events involve firms that engage in frequent patent litigation. Hence, we construct trimmed samples to exclude all events whose event windows overlap with other events. After trimming, we have 7,798 alleged infringer events.[23] We refer to the groups of observations outside the trimmed samples as the overlap samples.

Table 3 reports the names of explanatory variables used in the cross-sectional analysis of alleged infringer events (Subsection 4.6). It is useful to discuss these variables here, because they highlight typical characteristics of observations that get trimmed out. Table 4 reports characteristics of these variables for alleged infringer events in the full sample, the 5-day trimmed sample and the 5-day overlap sample. The far right column reports a z-statistic from a test of the difference in the means of the trimmed and overlap samples.

Because we lack strong theoretical predictions or priors about the relationship between firm size and CARs, we specify the $size_1 - size_5$ dummy variables for revenue categories. The $size_1$ firms are the smallest and the $size_5$ firms are the largest. Categories are defined in $2010 dollars, and for revenue the year of the event.

Generally, observations in the overlap sample include firms that are less likely to be newly public (1.5% vs. 4.1%) and are larger (average revenue $39.7 billion vs. $9.6 billion—see the Note to Table 4). Firms in overlap observations are also more likely to be sued by an NPE (49.8% vs. 38.6%), less likely to be a single defendant (39.3% vs. 50.4%), and are far less likely to be surprised by the lawsuit (3.5% vs. 33.4%). The overlap sample includes relatively more chemical/pharmaceutical and technology (computer, electrical/electronic, software and telecommunications) firms. The trimmed sample includes relatively more instrument, mis-

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[23]Events where event windows overlap with estimation window could also introduce additional noise and lead to inefficient estimates of the parameters in (3). In previous versions of this paper, our trimmed sample excluded such "estimation window overlaps" (see our online Appendix for additional details). However, subsequent analysis has shown that there is no evidence of efficiency loss in our data. For events with at least one estimation window overlap, the average variance for $\widehat{\psi}$ is .0038 (N=4,894), which is *lower* than for events with no estimation window overlaps (.0064, N=3,713). In our estimations of average CARs and in our cross-sectional analysis, we weight all CARs using the inverse variance of the estimate of ψ as the weight.

cellaneous manufacturing, and retail/wholesale firms.

4.5. Estimation of CARs and Average CARs

Following equation (3), we estimate a CAR for each event using a 5-day window. To estimate the average CAR of patent litigation filing, we use a weighted mean of all estimated abnormal returns, where the weight for each observation is the inverse of the variance of the estimate of ψ for that observation. The limiting distribution of this efficient test statistic is asymptotically normal (Salinger 1992). The null hypothesis is that the average CAR is zero.

Table 5 reports estimated average and median CARs (for both windows) for the full sample, trimmed samples and overlap samples. For the full and trimmed samples, the estimated average CAR is negative and significant when all events are included and when just alleged infringer events are included, but is insignificant when just patentee events are included. Hence, in the context of our theoretical model, we conclude that $\Delta V^R < 0$, but cannot reject the hypothesis that $\Delta V^P = 0$. Based on these results, we focus on alleged infringer events. We also use Proposition 2's prediction that if the probability of settlement is sufficiently high, then the effect of litigation on alleged infringers (ΔV^R) understates litigation costs.

For the trimmed sample, we estimate that alleged infringers sued for patent infringement lose -0.18% of their value on average. The estimate is highly statistically significant. The estimated median is of higher magnitude, indicating a skewed distribution. For the overlap sample, the estimated average CAR is -0.25%, higher (in absolute value) than for the trimmed sample. This suggests possible double counting of losses. By omitting these observations from our aggregate cost estimates, we reduce the (absolute value of the) average CAR from -0.19% (full sample) to a more conservative -0.18%.[24]

Our estimate of average percent losses are well below typical 2-3% estimates relying on announcements in the *Wall Street Journal* or similar periodicals (Bhagat, Brickley and Coles

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[24]We also estimate average CARs without applying weights to the individual CARs. Statistical significance remains the same for the full and trimmed samples. Estimated average CARs are a bit higher in magnitude for all events and for alleged infringer events. Hence, our use of weights in the main analysis produces a more conservative estimate of costs. The estimated average CAR for alleged infringers for the overlap sample is insignificant without weights.

1994; Bhagat, Bizjak and Coles 1998; Lerner 1995). Our estimates are lower (in absolute value) than the average CAR estimated by Bessen and Meurer (2008, 2012), which was around -0.50%. Part of the reason is our use of a 5-day window, which is more appropriate for these data and produces more conservative estimates of average CARs and aggregate costs.[25]

4.6. Cross-Sectional Variation in CARs

We now estimate a cross-sectional model of event CARs. Our preferred specification relates firm size, surprises and CARs:

$$CAR_i = \beta_1 + \beta_2 size2_i + \beta_3 size3_i + \beta_4 size4_i + \beta_5 size5_i + \beta_6 surprise_i + \epsilon_i, \qquad (4)$$

where CAR_i is the 5-day cumulative abnormal return for event i and ϵ_i is an error. We estimate this model using weighted Ordinary Least Squares (OLS) regression, with the inverse of the variance of the estimated ψ [from (3)] as the weight.

The first column of Table 6 shows results for the 5-day trimmed sample. Observations during 1984 and 1985 are omitted due to our inability to classify the surprise variable, so the sample sizes are smaller than in Table 5. Variables are scaled so that the coefficients are interpreted as percentage-sized effects. Since $size_1$ is omitted, these results indicate that firms of the smallest size involved in non-surprise lawsuits lose about 1.07% of value. Larger firms lose significantly less, but the relationship between size and losses is not clearly a monotonic function of firm size. The coefficient estimates indicate that firms in the $size_2$ category see estimated gains of about 0.05% (non-surprise lawsuits), while the largest ($size_5$) firms lose about 0.14% (non-surprise lawsuits). Given that the $size_2 - size_5$ coefficients are so similar, however, we cannot clearly distinguish between the effects across these groups. If the case is a surprise, firm value falls by an additional 0.21%, which is significant at the 10% level.

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[25]See our appendix for details. We report additional estimates using a 25-day window, which allows for more direct comparison of our results with Bessen and Meurer (2008, 2012) and for illustration of the implications of using the 5-day window for our main estimates.

We then estimate the model including $newfirm$, npe_case and $single$, as well as industry and year dummies (column 2). The coefficient estimates for the variables in the preferred specification do not change demonstrably. None of the coefficients on the three added variables are statistically significant. The industry dummies are also not jointly significant at the 10% level. Perhaps most surprisingly, the effect of npe_case is negative but insignificant. Hence, alleged infringers in NPE cases fare no differently than in other cases. This is robust to specifications that omit other variables (columns 3-4), so this result is not driven by firm size or lawsuit surprises. When we estimate the model using the 5-day full sample (column 5), the basic pattern holds but $size5$ firms have higher losses (-0.22%, non-surprise lawsuits).

Our finding that the smallest firms suffer the largest percentage losses is consistent with the conjecture that business risk is an important driver of costs. Namely, if business risk is important, then the harm that the litigation causes for the day-to-day profitability of a firm's business should fall with firm size. Small firms tend to have a higher percentage of their business in a lower number of product lines (Gollop and Monahan 1991), so that when litigation causes the firms to delay innovating on these product lines, they will lose relatively more than larger firms. And for very small firms, a single patent lawsuit could destroy it.

The possible non-monotonic effects of size, though small when compared to the effect of $size1$, are somewhat puzzling. For larger firms, product lines are more diversified and patent disputes are more of a routine part of doing business, so business risks of litigation should be smaller. However, for the very largest firms there will typically be a lot at stake in disputes. In such circumstances, discovery costs and other litigation costs will rise, and this could offset the business risk effect.[26]

The effect of *surprise* in specifications (1) and (5) implies that, when the firm has not recently been involved in a patent lawsuit, the loss more than doubles for all but the smallest firms. It is hard to imagine how such a doubling could occur as a result of direct legal costs. More likely, it reflects a greater tendency for firms surprised by litigation to make strategic errors in adjusting innovation and production decisions to a litigation environment. Clearly,

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[26]Informally, patent attorneys have also indicated to us that (in their experience) disputes involving big, directly-competing firms (e.g., Polaroid-Kodak from the 1970s-80s, or Apple-Samsung more recently) are more likely to proceed to trial. This conjecture is an interesting topic for future research.

this effect is not as robust as the effects of size. When the additional variables and year and industry dummies are included in specification (2), the coefficient is not significant at the 10% level. We interpret the size of this effect with more caution.

This effect could also reflect different probabilities of litigation. However, for this to explain a doubling in costs, λ (from the section 3 model) must be extremely high in the non-surprise cases. If a surprise lawsuit is perceived as a zero probability event, then λ must be at least .5 in non-surprise lawsuits to generate the effect. Finally, if this effect does explain part of the variation in CARs, then it also causes us to undercount costs.

4.7. Estimating Aggregate Costs

We estimate a cost for each event and then aggregate them. We consider three different approaches to estimating event costs. Under the *Average CAR Method*, we multiply the average CAR from Table 5 by the market value of each event party. The *Individual CAR Method* multiplies the CAR for each event by the event party's market value. The *Linear Model Method* uses the model from (4) and the parameter estimates from Table 6, column 1, to predict CARs. Then, we multiply the estimated CAR by the event party's market value. We report estimates of total losses, for the trimmed and full samples, in Table 7. When some observations are trimmed out, we also report a "scaled up" estimate of average cost that assigns to omitted observations the same *average dollar loss* as estimated in the sample.

The Average CAR method, used previously by Bessen and Meurer (2008; 2012), is simple to use, easy to understand, and imparts to each event the average effect of litigation. The third feature means that it lends itself to a sensible way to estimate costs intertemporally. For example, this method never estimates that alleged infringers lose from patent litigation in some years but gain in other years. Its disadvantage is that if CARs are driven by firm or event characteristics, then it may lead to biased estimates.

Under the Average CAR Method, we estimate an average loss of $46.2 million per event. Due to the skewness in firm valuation, this is considerably higher than the estimated $5.8 million median loss. Summing losses, we estimate about $361 billion in losses for the 7,798 observations. Scaling up to include the other 809 observations, we estimate about $398

billion in aggregate losses. For the full sample (average CAR = -0.31%), the estimated average loss is $53.4 million per event and the aggregate cost estimate is $460 billion.

The Individual CAR Method helps overcome the potential effects of cross-sectional variation, and with minimal parametric specification. However, it introduces significant noise. Estimated mean costs are more than twice as high as previous estimates. For the trimmed sample, the average event costs $115 million and the aggregate estimate is about $990 billion. Aggregating to the yearly level, we estimate seven years during 1984-99 where alleged infringers *gained* value due to being sued for patent infringement.

The Linear Model Method also helps to overcome the potential effects of cross-sectional variation, but it requires parametric choices. We estimate (4), then use $\widehat{\beta}$ to estimate losses by using the model to "predict" CARs for each event:

$$\widehat{CAR}_i = X_i\widehat{\beta}. \tag{5}$$

To predict the loss, we multiply this estimate times the value of the common stock of the firm at the time of the event. Mean costs are $41.4 million per event, slightly lower than under the Average CAR Method. The implied aggregate costs are about $308 billion unscaled and $356 billion scaled.

Hence, the Linear Model Method estimates the lowest costs. Estimates from the Average CAR Method are slightly higher, while estimates from the Individual Cost method are more than twice as high. Given our goal of estimating a lower bound of aggregate costs, we rely on event-cost estimates from the Linear Model Method.[27]

Prior estimates of external legal fees are typically well below our $41.4 million average cost per event. For example, the American Intellectual Property Law Association's 2011 survey estimates between $0.5 and $3.6 million in costs through discovery and between $0.9 and $6.0 million through trial (AIPLA 2011). Accounting for the fact that most patent lawsuits do not complete the discovery phase, Bessen and Meurer (2014) use the AIPLA

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[27]When we use just events where we can identify the alleged infringer for sure, we lose 1,530 events (about 20%). The average CAR for the remaining events is -0.20%, which is a bit higher in magnitude than the -0.18% estimate in Table 5. The coefficient estimates in the regressions are very similar to those in Table 6. The average loss using the linear model method is $43.3 million, also similar to Table 7. For patentee events, we lose 1,207 events. The average CAR remains not precisely identified.

numbers to estimate an average legal cost of \$483,000 per filed lawsuit. These estimates reflect a broad cross-section of litigating firms, including the very smallest.

Bessen and Meurer (2014) also carefully study both external legal fees *and* licensing costs for a sample of 82 firms involved in 1,184 defenses versus NPEs between 2005-11. They estimate a mean direct cost of \$7.91 million per suit. Legal costs comprise \$1.38 million and licensing costs comprise \$6.53 million.

Among alleged infringers in NPE cases during 2005-09, our Linear Model Method estimates the average cost to be \$53.4 million.[28] Using the Bessen-Meurer figures as benchmarks for direct litigation costs, these results imply that in our data the alleged infringer's indirect costs are just above \$45 million on average. Hence, about 85% of the overall costs are indirect.

Additional outside evidence suggests that indirect costs may indeed be extremely high. Tucker (2014) studies the response, to lawsuits filed by Acacia Technologies, of firms producing medical imaging technology. Sued firms *cease* incremental innovation in imaging IT—that is, they produce zero new variations of their existing products during the lawsuits. Relative to firms producing products not covered by Acacia's patents, the sued firms suffer a one-third reduction in demand for their technology.

The characteristics of this disruption are consistent with our estimates of indirect costs. Consider the following extrapolation off of Tucker's results. Suppose profit also falls by one-third for just one year of litigation ending in settlement, but then returns to its pre-litigation level and remains constant forever. Then the implied value of ωL_I as a percentage of firm value is (for an interest rate of 5%) about 1.6%.[29] This is similar to our Linear Model Method's predicted CAR for small firms. If the patent lawsuit affects about 10% of the alleged infringer's business, then the effect on the firm's value would be a -0.16% change, similar to our estimates of the average CAR for larger firms.

Related research shows that the effect on innovation goes beyond medical imaging. In a

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[28]These numbers reflect the event cost estimates under the Linear Model Method approach. We add up those events with npe_case=1. We also estimate an average loss of \$38.1 million per alleged infringer in a PE case during 2005-09 and \$47.4 million per alleged infringer overall for this period.

[29]If firms discount future years at rate $\frac{1}{1+r}$, then the percentage loss in present discounted value from a first-year, one-time drop in profit of $\frac{1}{3}$ is $\frac{1-\frac{1}{1+r}}{3}$, which is about 1.6% for $r = 0.05$.

study of a cross-section of firms sued by NPEs, Cohen, Gurun and Kominers (2014) show that those firms that succeed in getting cases dismissed spend an average of $211 million more on R&D and produce an average of about 64 more patents, compared to firms that either settled litigation or proceeded to trial. One prominent, but understudied, reason for this business risk is that sued firms that do not curtail releasing related new products may face claims of "willfully" infringing the technology in the patents asserted against them.

Importantly for our study, indirect costs of patent litigation are generally uninsurable. During our sample period, the market for patent litigation insurance is very small and typically covers direct litigation costs and damages only.[30] Moreover, policies typically schedule products or patents for coverage, and are therefore incomplete.

In principle, our technique can be used to study differences in indirect costs among subgroups, e.g, for non-troll lawsuits or by industry. Unfortunately, we do not have access to data on direct costs for these other categories. We can use the Bessen and Meurer (2014) estimates of direct costs to learn a bit more about differences in costs from NPE lawsuits by firm size. For small and medium-sized firms (less than $1 billion in annual revenue), they estimate a mean direct cost of $1.75 million while we estimate a mean total cost of $1.89 million. For large firms, they estimate a mean direct cost of $8.79 million while we estimate a mean total cost of $64.47 million. Hence, roughly 8% of costs are indirect for small and medium-sized firms, while 86% of costs are indirect for large firms. We regard these results as a bit speculative, given that the sets of lawsuits are sampled quite differently. But the large difference in the percent of costs that are indirect (by firm size) suggests that there could be substantial variation across types of cases in general, and motivate this as an important topic for future research.

5. Estimation of Rents

Researchers have long known how to use firm market value to estimate patent value. Griliches (1981) treats firm value as decomposable into component parts, including a part

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[30]This market has grown in recent years. See Llobet and Suarez (2012) and Lemus, Temnyalov and Turner (2017) for discussion of some key issues.

attributable to patent rents, and performs "hedonic" regressions of stock market value (usually Tobin's Q) on patent-related variables. Investor behavior—i.e. stock market value of patent-holding companies—reveals patent value.

Hayashi (1982) specifies the relationship between Tobin's Q and the value of rents for firms with market power in the following way:

$$Q_{jt} = \frac{V_{jt}}{K_{jt}} = q_t \left(1 + \frac{W_{jt}}{K_{jt}} \right), \tag{6}$$

where V_{jt} is firm $j's$ market value at time t, K_{jt} is its capital stock, W_{jt} is the present value of its rents and q_t is "marginal q," which can account for short-term disequilibrium in capital markets. Rents come from patents and from other features of technical knowledge:

$$W_{jt} = uP_{jt} + \mu_j K_{jt},$$

where P_{jt} is the number of patents obtained, u is the mean rent per patent and μ_j is the markup for rents earned through other means. The aggregate capital stock includes both tangible assets A_{jt} and knowledge. Following Hall (1993), K_{jt} is a weighted sum:

$$K_{jt} = A_{jt} + s_{jt} R_{jt},$$

where R_{jt} is dollars of R&D and s_{jt} indexes the "success" of R&D in generating valuable patentable innovations. Generally, one cannot observe s_{jt}. However, patent counts may serve as a proxy for it (Griliches 1981). The firm patents inventions where the value of rents with the patent minus the cost of patenting exceeds the rents that could be obtained without the patent. This idea is represented by a patent-propensity equation:

$$P_{jt} = R_{jt} G \left(s_{jt} \right),$$

where G is increasing in s_{jt}. If G is linear in s_{jt}, we have

$$s_{jt} = \alpha + \phi \frac{P_{jt}}{R_{jt}}.$$

Substituting, we have an equation relating the "traditional" measure of Tobin's Q, $\frac{V_{jt}}{A_{jt}}$, to R&D and patent stocks:

$$ln\frac{V_{jt}}{A_{jt}} = ln(q_t) + ln(1 + \mu_j) + ln\left(1 + \alpha\frac{R_{jt}}{A_{jt}} + \gamma_{US}\frac{P_{jt}}{A_{jt}}\right) + \epsilon_{jt}, \qquad (7)$$

where $\gamma_{US} = \frac{u}{1+\mu_j} + \phi$. We include year dummies to pin down $ln(q_t)$. If the fixed effects are ignored ($\mu_j = 0$), then this equation can be estimated using non-linear least squares (NLLS). Unfortunately, the estimate of γ will not converge to u :

$$plim \ \widehat{\gamma} = u + \phi.$$

This reflects the fact that patents affect the relationship between rents and Tobin's Q in two different ways—through direct contribution to rents and through correlation with R&D quality and success. Assuming $\phi > 0$, we see that $\widehat{\gamma}$ does converge to an upper bound for u.

We also consider a specification where firm value depends on patents from other countries. Denoting γ_{US} as the upper bound for rents on US patents and γ_{EPO} as the upper bound for rents on EPO patents and again ignoring fixed effects, we write:

$$ln\frac{V_{jt}}{A_{jt}} = ln(q_t) + ln\left(1 + \alpha\frac{R_{jt}}{A_{jt}} + \gamma_{US}\frac{P_{jt}}{A_{jt}} + \gamma_{EPO}\frac{P_{jt}^{EPO}}{A_{jt}}\right) + \epsilon_{jt}. \qquad (8)$$

Importantly, this specification helps us avoid some double-counting of rents. While the underlying invention is the same for an EPO & US patent, we want to measure the extent to which the patent creates market value beyond the invention. The patents for each jurisdiction provide market exclusion in each jurisdiction and this is the ultimate source of rents. When we don't control for EPO patents, the US patent data proxies for the EPO value because the two are correlated. Indeed, we expect the coefficient we estimate for US patents in (7) to be inflated because it includes both rents from the US patent and, depending on the correlation coefficient, some portion of rents from the EPO patent. Hence, the additional variable in (8) produces a more accurate estimate of the US patent rents.

For estimating US rents, the inclusion of European patents does not introduce a sample selection bias. The sample of firms is the same with and without European patents. For

estimating EPO rents, however, our sample is highly selected and we will not interpret γ_{EPO} as the average rent for all EPO patents. We discuss this further in subsection 5.2.

5.1. The Data

The data we use to estimate patent rents span 1969-2006.[31] Patent data come from the USPTO, EPOLINE and PATSTAT. Financial data, used to construct variables for firm value, assets and R&D, come from COMPUSTAT. Our USPTO patent variables come from the updated NBER data originally captured by Hall et al. (2001). These data include a virtually comprehensive collection of USPTO variables for 1976-06. The updated NBER data also include an improved match (using the CUSIP identifier) to publicly-traded firms' financial data (including R&D expenditures) stored in COMPUSTAT (Bessen 2009a), and this increases the number of matches over the original NBER data. COMPUSTAT firms are also matched to EPO patents using the same algorithm. The EPO patent data come from EPOLINE and PATSTAT, online archives of EPO patent documents.[32]

We identify all firms publicly traded in the United States that match to at least one patent during 1979-2006. We then exclude any firm that performs less than \$2,000 of R&D over a 3-year period, and exclude firms without 4 years of non-missing data on key variables. This yields 4,481 firms. We also use data from 1969-78 to construct patent (and R&D) stocks, but do not use market-value variables from this period. Our main regressions use all variables from 1979-2002. Finally, because we use variables that are ratios, variables in the tails of the distribution introduce significant measurement error. We trim the highest and lowest 1% of the $\frac{V}{A}$ observations that we use to estimate equations (7) and (8).

Sample statistics are shown in Table 8. We have 40,287 observations from 1979-2002. Of these, 29,935 come from 1979-1997, nearly 16% more than Bessen (2009b) uses for the same period. Nearly all of this increase is due to the improved matching routine.

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[31]See the online appendix for supporting details on data sets and variable construction. In addition to those referenced below, we rely on techniques introduced by Brainerd et al. (1980), Lewellen and Badrinath (1997) and Thoma et al. (2010).

[32]The data can be downloaded at http://www.researchoninnovation.org. Thanks to Grid Thoma for executing the matching.

Firm market value, V, is the sum of the value of all claims on the firm, including the common stock, preferred stock, long-term debt adjusted for inflation, and short-term debt net of current assets. The value of assets, A, is the sum of the net book value of plant, equipment, inventories, accounting intangibles, and investments in unconsolidated subsidiaries, all adjusted for inflation. The calculation of the R&D stock follows Hall (1990). For each of these variables, the mean and median are higher, but by less than 25%, than in Bessen (2009b).[33] Part of this difference is due to additional variables found due to improved matching, and part due to the added 1998-2002 data.

The patent stock is based on the patent applications held by a firm during a particular year. Applications command significant option value and investors update expectations of future rents before the patent offices complete their examination processes. The drawback to application-based stocks is that issue dates lag application dates, so patent data through 2006 permit accurate application counts through only 2002. We use the NBER data to construct US patent counts, and use EPOLINE and PATSTAT to construct European counts. Patents are depreciated 15% each year. To interpret the γ coefficient in constant 2010 dollars, we scale the ratio $\frac{P}{A}$ by the GDP deflator. The mean US patent stock here, 106.7, is about 20% higher than in Bessen (2009b), while the median patent stock is nearly identical. The mean EPO patent stock is ... Tobin's Q is also bit higher, on average (1.00 vs. 0.85) and at the median (0.71 vs. 0.61), in our paper.

5.2. Estimation

We first estimate the model in (7) and report results in column 1 of Table 9. The coefficient estimate for γ implies an upper bound patent rent of $517,000. It also implies that an upper bound for the aggregate, present-discounted value of the flow of patent rents, at a given moment in time, is the total size of the patent stock times $517,000. In real terms, this figure is a bit smaller than the same estimate reported by Bessen (2009b), reproduced in column 5. The estimate of α is statistically indistinguishable from 1. Column 3 shows

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[33]To get an apples-to-apples comparison, convert our statistics back to 1992 dollars using the GDP deflator of about 1.449. The mean V here of $2,815.32 is about $1,942.72 in 1992 dollars. This is about 24% higher than the mean V of $1,568 in Bessen (2009b).

estimates of (7) with just 1979-97 data. The estimate is a bit higher, suggesting the possibility that private rents fell during 1998-2002.

Results from a regression of the model in (8) are shown in column 2 of Table 9. The coefficient estimates in column 1 imply an upper bound rent of \$351,000 for US patents. Compared to the estimate from column 1, this reduces the estimated average rent by 32%. By controlling, somewhat, for patenting outside the US, we estimate a more precise upper bound for US rents.

The estimate for EPO patents, \$1,821,000, is strikingly high. While this is consistent with other research in suggesting that patents taken out in both Europe and the US are high (Hall, Thoma and Torrisi 2007), we do not interpret this as an average rent for EPO patents. Among firms holding EPO patents, our group of firms is highly selected and omits lower-valued firms (which will tend to hold lower-valued EPO patents). Moreover, it would be wrong to compare this figure to the \$351,000 figure for US patent rents. In essence, the endogeneity problem (of successful R&D being correlated with firm value) is likely more severe for EPO patents. For companies publicly-traded in the United States, filing patents in the EPO (or other foreign offices) should be more likely for inventions that they already expect to have a high market potential. If high-value firms tend to own more EPO patents, the upward endogeneity bias in our estimates of EPO rents is likely to be quite high.

That said, there are some reasons why EPO patents might be quite valuable, First, it is clear that patents are useful for blocking rivals (Cohen et al 2000; Blind et al 2006; Blind et al 2009), and EPO patent applications are particularly effective because they block the whole European market.[34] Second, quality standards at the EPO may be higher (van Pottelsberghe de la Potterie 2011). If patents must clear a higher bar, they will tend to be worth more. Third, the EPO patent stock proxies for patenting in multiple jurisdictions. Because an EPO application is much more expensive than a national application, it only

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[34]Patent applicants at the EPO have to designate the targeting states sixth months after the application has been published, implying that they have quite a long time to choose in which markets they seek actual patent protection. This can even be prolonged until a patent is granted, since applicants can withdraw designated states up to the grant data and the average grant lag at the EPO is 5-7 years. In addition, the PCT route can be used which gives an applicant additional time to decide what to do with a patent even before it enters the EPO system. A PCT application further comes with a preliminary search report on patentability, which delivers additional information to the applicant. See Frietsch, Neuhäusler and Rothengatter (2012).

makes economic sense to file an EPO application when targeting more than one country.

Here, the most we can say is that R&D-performing, publicly-traded firms in the United States that obtain EPO patents are worth considerably more than other such firms. Carefully disentangling rents from signaling effects is a challenging topic for future research. We look forward to more careful comparisons of US and EPO patent rents.

6. Comparing Private Costs and Private Rents Across Time

Recall Figure 1 (p. 4). To estimate an upper bound of aggregate rents by year, we follow Bessen and Meurer (2008) and assume an annual flow profit of 15% of the stock of patents. We multiply the average rent from Table 3, $351,000, by 15% and by the size of the aggregate patent stock for each year. Our estimates do not include 2003-2009 because of truncation. Our lower-bound aggregate cost estimates rely on event losses using the Linear Model Method on the 5-day trimmed sample.

We have cost and rent estimates for 1986-2002. During this period, private costs sum to about $114 billion and private rents sum to about $182 billion. Hence, the upper bound of private rents exceeds the lower bound of private costs by about 60%. However, by the late 1990s the faster-trending private costs had caught up with private rents, exceeding them during 1999-2000 and falling below but very close to them in 2001-02.

Assuming average (1984-2002) growth of 6.2% in private rents during 2003-09, we estimate an upper bound of about $159 billion in rents for these years.[35] We estimate a lower bound of about $194 billion in private costs. Estimated costs exceed estimated rents for all years during 2005-09. For the 1986-2009 period, we estimate rents of about $342 billion and costs of about $308 billion.

We also construct costs of PE cases by including all events such that $npe\_case = 0$. Upper bound private rents meet or exceed this part of private costs during all years. Growth in these costs after 2000 roughly mirrors growth for *total* private costs prior to the mid-1990s.

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[35]Measuring growth rates since 1984 removes patenting and litigation activity prior to the establishment of the CAFC, which many researchers argue prompted an increase in patenting. If 1979-84 data are included, then the average annual growth in rents is actually below 5%.

Importantly, we use the same matching algorithm for both parts of the analysis (Bessen 2009a). Hence, for firms that we miss rents, we also miss costs. And because we trim out some litigation events due to overlaps, we miss relatively more costs.

7. Discussion

We find that patent litigation costs exceed patent rents during 1999-2000 under the most conservative comparison of costs and rents. Moreover, the trend in litigation costs suggests that unless private rents rose sharply during the mid-to-late 2000s, private costs of patents exceed private rents of patents for 2005-2009 as well. The main driver is that litigation costs grow more than twice as fast as rents. Part of this growth appears to come from the emergence of NPE cases in the mid-to-late 1990s.

From a policy perspective, these results suggest that the patent notice problem identified by Bessen and Meurer (2008) has gotten worse since 1999. In emphasizing the importance of indirect costs, we also highlight the potential effects of the doctrine of willful infringement. Tucker (2014) and Cohen, Gurun and Kominers (2014) both cite fear of willfulness claims, and ensuing trebling of damages, as a primary reason that firms curtail innovation and new product introduction. Our findings that indirect costs comprise around 85% of total patent litigation costs motivate careful study of the efficacy of the willfulness doctrine.[36]

One potentially fruitful direction is to study the effects of two important recent decisions affecting the willful infringement standard. The CAFC's 2007 *Seagate* decision imposed a two-part test that effectively made it more difficult for a patentee to prove willfulness,[37] while the Supreme Court's 2016 *Halo* decision repudiated this test.[38] In our data (recall Figure 1), estimated litigation costs are lower in 2008 and 2009 than in 2007. With just the yearly aggregates from Figure 1, it is hard to disentangle a potential Seagate effect from other cases (e.g., *eBay v. MercExchange*) that might also affect the liability that alleged infringers face. It would be interesting to study richer and more recent data to see if these

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[36]Researchers have highlighted other inefficiencies in the willfulness doctrine—namely, that its uncertain applicability leads to wasted resources in litigation (see, e.g., Means 2013).

[37]*In re Seagate Technology LLC,* 497 F.3d 1360 (Fed. Cir. 2007) (en banc).

[38]Halo Electronics, Inc. v. Pulse Electronics Inc., et al., 579 US 1 (S.C. 2016).

2007 and 2016 events affect costs faced by alleged infringers.

Our interpretation of our statistical results is subject to a few possible caveats. First, we view the lost wealth of investors as the direct and indirect costs due to litigation. Investors bid the stock lower because they expect lower future profits, including the effects of resolving the suit. If the efficient market hypothesis fails to hold, then our interpretation is incorrect. Perhaps investors don't account for possible successful resolutions and bid the stock back up after the suit is resolved. However, outside evidence (Bessen and Meurer 2012, pp. 76-77) finds that such losses are permanent.[39]

Second, the observed loss in firm value might reflect new information that the lawsuit reveals about management quality, shirking, competitors' entry plans, or technology quality. Such information might lead investors to lower their profit expectations. Bessen and Meurer (2012, pp. 83-87) perform additional checks on these considerations, and conclude that investors' estimates of losses due to litigation are at least as great as the estimated CARs.

Third, patents may provide positive innovation incentives even if costs exceed rents in aggregate. For instance, suppose some firms innovate and patent and a separate group of firms copy. Then innovator firms are the only agents for whom good innovation incentives (from patents) are important. If aggregate costs exceed aggregate rents, the innovators might still have positive rents while copiers have negative rents; then, the patent system would promote innovation. However, Bessen and Meurer (2013) show that firms that invest more in R&D are significantly more likely to be sued, all else equal. Hence, innovators bear the costs of litigation, making a comparison of aggregate costs and rents appropriate.

Fourth, our paper considers just publicly-traded firms, so the conclusions regarding incentives apply only to this group. If non-public firms, such as NPEs, strictly gain from monetization through lawsuits *and* public patentees gain from these lawsuits through ownership rights, then we would miss some public-firm rents. While there have been some recent cases of large companies forming private "hybrid patent-assertion entities" to gain strategic

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[39]In an important objection to the use of event studies to measure changes in wealth, Lunney (2008) notes examples where stock-market reactions to news appear to be irrationally large, e.g. Eli Lilly's large losses after an adverse August 9, 2000 ruling on one of its patents covering Prozac (Lunney, 2008, pp. 44-50). Our interest is in the mean effect of a large number of lawsuit filings, and the inclusion of a large number of such events allows us to discern the signal despite the noise.

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advantages (Scott Morton and Shapiro 2013), there is little evidence of such NPEs in our 1984-2009 data. In addition, Bessen, Ford and Meurer (2011) find little evidence that damages collected by NPEs go to inventors. And NPEs often target small non-public firms, so any NPE gains would need to be balanced against losses by non-public firms in adjusting our cost-benefit comparison. But given our estimates of no lawsuit effects for patentees and strictly negative lawsuit effects for alleged infringers (among public firms), we find it unlikely that rents to NPEs (or other non-public firms) from suing for patent infringement outweigh costs to non-public firms defending against such suits.

Finally, patents may sometimes be used for "defensive" purposes, as bargaining chips in settling litigation. Such uses make it harder to think about separating rents from costs, because a firm may realize value from its patents as savings on litigation costs or damages. But patents have defensive value only in cases where they may be asserted for something significant against a firm that might similarly assert patents against them. That is, patents have defensive value only if they may earn rents (Lemley and Melamed 2014). In our model, when firms efficiently settle litigation they may execute agreements that sacrifice these rents in exchange for access to other firms' patents. But the total rents achieved by the firms in the bargain still rises. So any litigation costs that we estimate using the event study will be undercounted due to this upward adjustment in joint rents. At the opposite extreme, if all firms were to farm out their patents to NPEs, then defensive patenting would not work at all. But the firms would still enjoy rents from their patents, and would capture them in the traditional way rather that through barter in patent litigation (Lemley and Melamed 2014).

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	All Cases		Public Cases	
	N	%	N	%
Case Type				
Not Identified	12,150	34.4	2,590	27.3
Identified	23,151	65.6	6,888	72.7
Infringement	19,259	83.2	5,911	85.8
Declaratory Judgment	3,892	16.8	977	14.2
Number of Defendants				
1	27,777	78.7	6,583	69.5
2 or more	7,523	21.3	2,895	30.5
Main Patent Type				
Not Identified	328	0.9	56	0.6
Identified	34,973	99.1	9,422	99.4
Design	2,790	8.0	430	4.6
Plant	21	0.1	0	0.0
Reissue	851	2.4	243	2.6
Utility	31,311	89.5	8,749	92.9
Main Patent Tech. Class				
Design/Plant	2,811	8.0	430	4.5
Not Identified	1,494	4.2	480	5.1
Identified Utility/Reissue	30,996	87.8	8,568	90.4
Chemical	3,068	9.9	889	10.4
Computers/Communications	5,862	18.9	2,359	27.5
Drugs/Medical	4,909	15.8	1,983	23.1
Electrical/Electronic	3,350	10.8	982	11.5
Mechanical	5,045	16.3	988	11.5
Other	8,762	28.3	1,367	16.0

Table 1: *Derwent Lawsuits, 1984-2009*

Note: These statistics reflect lawsuits recorded by *Derwent Litalert* for years 1984-2009. Each Derwent record includes the name(s) of at least one plaintiff and at least one defendant, as well as a "main" patent. We match these names to public firms using an algorithm developed by Jim Bessen. For the categories that are indented, the statistics are conditional probabilities. For example, the probability a *Derwent* case is an infringement suit, conditional on being identified as either an infringement suit or a declaratory judgment, is 83.2%. We use the "main" patent to construct statistics on patent type and technology class. All calculations are performed in STATA.

	Derwent		Non-Derwent	
	N	$\%$	N	$\%$
Case Type				
Not Identified	0	0.0	0	0.0
Identified	954	100.0	460	100.0
Infringement	952	99.8	460	100.0
Declaratory Judgment	2	0.2	0	0.0
Number of Defendants				
1	405	42.5	230	50.0
2 or more	549	67.5	230	50.0
Main Patent Type				
Not Identified	1	0.1		
Identified	953	99.9		
Design	9	0.9		
Plant	0	0.0		
Reissue	16	1.7		
Utility	928	97.4		
Main Patent Tech. Class				
Design/Plant	9	0.9		
Not Identified	76	8.0		
Identified Utility/Reissue	869	91.1		
Chemical	6	0.7		
Computers/Communications	670	77.1		
Drugs/Medical	8	0.9		
Electrical/Electronic	93	10.7		
Mechanical	56	6.4		
Other	36	4.1		

Table 2: *Patent Freedom Lawsuits, 1990-2009*

Note: These statistics reflect lawsuits captured by *Patent Freedom* for 1990-2009. Some of these cases match to *Derwent* records, while others do not. For the categories that are indented, the statistics are conditional probabilities. For example, the probability a *Patent Freedom* case (which matches a *Derwent* record) is an infringement suit, conditional on being identified as either an infringement suit or a declaratory judgment, is 99.8%. The *Patent Freedom* data do not include patent numbers, so it is not possible to construct statistics about patent type or class for those observations which do not match to a *Derwent* record. All calculations are performed in STATA.

Variable	Type	Description
newfirm	Dummy	The firm has been public less than 5 years.
npe_case	Dummy	"1" if the lawsuit includes an NPE.
revenue	$USm (2010)	The party's previous-year revenue.
single	Dummy	"1" if the party is the only defendant.
size1	Dummy	"1" if revenue <$100m.
size2	Dummy	"1" if $100m<revenue<$500m.
size3	Dummy	"1" if $500m<revenue<$5b.
size4	Dummy	"1" if $5b<revenue<$20b.
size5	Dummy	"1" if revenue>$20b.
surprise	Dummy	"1" if party is in a lawsuit in the past 2 years.

Industry Categories

Variable	Type	Description
chem & pharm	Dummy	"1" if the firm is in SIC 28.
computers	Dummy	"1" if the firm is in SIC 35.
electrical & electronics	Dummy	"1" if the firm is in SIC 36.
instruments	Dummy	"1" if the firm is in SIC 38.
misc manufacturing	Dummy	"1" if the firm is in SIC 20-27, 29-34, 37, 39.
retail & wholesale	Dummy	"1" if the firm is in SIC 50-65, 67.
software	Dummy	"1" if the firm is in SIC 73.
telecommunications	Dummy	"1" if the firm is in SIC 48.
other	Dummy	"1" if the firm has any other SIC code.

Table 3: *Variables for Cross-Sectional Analysis*

Note: SIC categories are obtained via CRSP.

Variable	(1) Full Sample			(2) Trimmed Sample			(3) Overlap Sample			DIM
	N	%	SD	N	%	SD	N	%	SD	z-stat
newfirm	8,442	3.9	13.7	7,642	4.1	19.8	801	1.5	12.2	5.35
npe_case	8,607	40.1	49.0	7,798	38.6	48.7	809	49.8	50.0	-6.08
single	8,607	49.3	49.9	7,798	50.4	50.0	809	39.3	48.8	6.11
size1	8,442	7.2	26.5	7,641	7.6	26.5	801	2.5	15.6	8.10
size2	8,442	12.0	33.2	7,641	12.9	33.5	801	3.6	18.6	12.21
size3	8,442	29.9	46.1	7,641	31.2	46.3	801	18.0	38.4	9.06
size4	8,442	22.7	41.3	7,641	22.8	42.0	801	21.7	41.2	0.72
size5	8,442	28.2	44.5	7,641	25.5	43.6	801	54.2	49.8	-15.69
surprise	8,369	30.5	46.1	7,566	33.4	47.2	803	3.5	18.4	35.37
chem & pharm	8,607	10.5	30.6	7,798	10.3	30.4	809	12.5	33.1	-1.81
computers	8,607	12.9	33.5	7,798	12.2	32.7	809	18.9	39.2	-4.70
electrical & electronics	8,607	16.8	37.4	7,798	16.3	36.9	809	20.8	40.6	-3.03
instruments	8,607	8.2	27.4	7,798	8.8	28.3	809	2.8	16.5	9.05
misc manufacturing	8,607	13.8	34.5	7,798	14.6	35.3	809	5.6	23.0	9.98
retail & wholesale	8,607	13.4	34.1	7,798	13.9	34.6	809	9.0	28.6	4.54
software	8,607	8.9	28.4	7,798	8.3	27.6	809	14.2	34.9	-4.66
telecommunications	8,607	4.9	21.6	7,798	4.7	21.2	809	24.7	43.1	-12.89
other	8,607	10.7	30.9	7,798	10.6	30.8	809	11.5	31.9	0.77

Table 4: *Descriptive Statistics, Alleged Infringer Litigation Events*

Note: These statistics reflect the characteristics of event parties in lawsuits during 1984-2009. Revenue data come from matching public firms to COMPUSTAT, while industry classes come from CRSP. Mean revenue is $23,253.0 million (N=8,442, SD=$43,708 million) for the Full Sample, $9,558.0 million (N=7,641, SD=$21,525.8 million) for the trimmed sample and $39,731.0 (N=801, SD=$47,730.2) for the overlap sample. In some cases, we could not identify a GVKEY identifier to obtain a certain match to COMPUSTAT. This is why the number of observations for the revenue-related variables is smaller. The number of observations for the *surprise* variable is smaller because we do not have enough data to capture this variable for lawsuits in 1984-85. The far right column reports the z-statistic from a difference-in-means (DIM) test of the mean of the trimmed sample minus the mean of the overlap sample. The DIM z-stat for mean revenue is -27.25. All calculations are performed in STATA.

	(1) Full Sample	(2) Trimmed Sample	(3) Overlap Sample
All Event Parties			
N	13,526	12,265	1,261
Mean	-0.11%*** (0.03%)	-0.10%*** (0.04%)	-0.08% (0.10%)
Median	-0.24%	-0.27%	0.02%
Alleged Infringers			
N	8,607	7,798	809
Mean	-0.19%*** (0.04%)	-0.18%*** (0.05%)	-0.25%** (0.13%)
Median	-0.32%	-0.34%	-0.04%
Patentees			
N	4,919	4,467	452
Mean	0.05% (0.06%)	0.03% (0.06%)	0.23% (0.16%)
Median	-0.12%	-0.14%	0.13%

Table 5: *Estimated Average Cumulative Abnormal Returns, 1984-2009*

Note: These statistics reflect an event study of patent litigation by all public firms matched to either a *Derwent* record or a *Patent Freedom* record during 1984-2009. We use equation (3). To estimate mean cumulative abnormal returns (CARs), we weight each individual CAR by the inverse of the variance of the estimated effect of the event. All calculations are performed in STATA. If two or more firms are involved in the same case, then each firm's participation is treated as a separate event. The "Patentees" category includes all firms clearly identified as patentees in known infringement suits or declaratory judgments, as well as plaintiffs in cases where we do not know if the case was an infringement suit or a declaratory judgment. The "Alleged Infringers" category includes all firms clearly identified as alleged infringers in known infringement suits or declaratory judgments, as well as defendants in cases where we do not know if the case was an infringement suit or a declaratory judgment.

	(1)	(2)	(3)	(4)	(5)
constant	-1.07**	-1.51***	-1.21***	-0.18***	-1.00***
	(0.47)	(0.54)	(0.45)	(0.62)	(0.45)
size2	1.12**	1.16**	1.15**		1.12**
	(0.50)	(0.51)	(0.50)		(0.49)
size3	0.91*	0.94**	0.97**		0.90**
	(0.47)	(0.48)	(0.46)		(0.45)
size4	0.95**	0.99**	1.06**		0.90**
	(0.47)	(0.48)	(0.46)		(0.46)
size5	0.93**	0.99**	1.08**		0.78*
	(0.47)	(0.49)	(0.46)		(0.46)
surprise	-0.21*	-0.15			-0.22*
	(0.12)	(0.13)			(0.12)
npe_case		-0.11	-0.07	-0.04	
		(0.12)	(0.10)	(0.010)	
newfirm		-0.15			
		(0.32)			
single		-0.05			
		(0.11)			
Industry and Year Dummies	No	Yes	No	No	No
Sample	Trimmed	Trimmed	Trimmed	Trimmed	Full
N	7,434	7,434	7,434	7,566	8,229
R^2	0.002	0.007	0.001	0.000	0.002

Table 6: *Cross-Sectional Analysis of CARs, 1986-2009*

Note: The results in column (1) reflect regressions using the equation $CAR_i = \beta_1 + \beta_2 size2_i + \beta_3 size3_i + \beta_4 size4_i + \beta_5 size5_i + \beta_6 surprise_i + \epsilon_i$, where CAR_i is the estimated cumulative abnormal return for a 5-day window from the event study described in Table 5. The equations are estimated using weighted ordinary least squares, where the weights are the inverses of the variances of the estimated effects of the events. All calculations are performed in STATA, and robust standard errors are in parentheses. Variables are scaled so that coefficient estimates are interpreted as percentage-point-sized effects. The following denote statistical significance: *** 1% level, ** 5% level, * 10% level.

	(1) Trimmed Sample	(2) Full Sample
Average CAR Method		
N_E (Estimation)	7,798	8,607
Mean	46.2	53.4
Median	5.8	7.0
Total Costs (N_E)	360,965.3	460,241.2
Total Costs (scaled up to N=8,607)	398,413.5	460,241.2
Individual CAR Method		
N_E (Estimation)	7,798	8,607
Mean	115.0	118.0
Median	1.6	1.6
Total Costs (N_E)	897,024.7	1,015,439.0
Total Costs (scaled up to N=8,607)	990,086.1	1,015,439.0
Linear Model Method		
N_E (Estimation)	7,434	8,229
Mean	41.4	61.8
Median	6.7	6.3
Total Costs (N_E)	308,174.1	508,865.7
Total Costs (scaled up to N=8,607)	356,800.4	532,240.5

Table 7: *Estimates of Aggregate Costs, 1984-2009*

Note: All results use the CARs from event studies performed according to the note in Table 5, using just alleged infringer events from the trimmed sample and full sample. For each of the three methods used to estimate aggregate costs, we estimate a loss for each event and then sum the losses. To estimate each event cost, the Average CAR Method multiplies the average CAR from Table 5 by the market value of the event party. In contrast, the Individual CAR Method multiplies the CAR for each event by the firm's market value. Finally, the Linear Model Method uses the model from (4) and the parameter estimates from Table 6, column 1, to predict CARs. Then, we multiply the predicted CAR by the event party's market value. All calculations are performed in STATA. The "scaled up to N=8,607" estimates calculate the average cost among all events in the trimmed sample. This cost is then imparted to all events not in the trimmed sample. Estimates are in $millions (2010).

	Mean	Median
Firm market value, V	$2,815.32	$103.85
Accounting assets, A	$1,410.18	$38.18
R&D stock, R	$366.24	$21.04
US Patent stock	106.71	4.41
EPO Patent stock	23.10	1.37
$\ln\left(\frac{V}{A}\right)$	1.00	0.71
Percent observations with no patents	20.79	

Table 8: *Sample Statistics*

Note: These statistics reflect 40,287 observations for 4,481 firms during 1979-2002. All calculations are performed in STATA.

Years	1979-2002	1979-2002	1979-1997	1979-1997	1979-1997 (Bessen 2009b)
γ_{US}	0.517*** (0.030)	0.351*** (0.029)	0.547*** (0.035)	0.381*** (0.035)	0.536*** (0.035)
γ_{EPO}		1.821*** (0.141)		2.012*** (0.168)	
α	1.031 (0.019)	0.997 (0.019)	1.183*** (0.025)	1.142*** (0.025)	0.992 (0.023)
N	40,287	40,287	29,935	29,935	25,681
Adjusted R^2	0.666	0.666	0.430	0.434	0.625

Table 9: *Estimates of US and EPO Patent Rents Using Application Stocks*

Note: The results in column 1 reflect non-linear least squares (NLLS) estimates of equation (7), while column 2 reports estimates of equation (8). The left-hand column lists the parameters estimated. All calculations are performed in STATA. Standard errors in parentheses are robust to heteroscedasticity. The estimates for γ are in millions of 2010 dollars. The null hypothesis for the first coefficient estimate is that the coefficient is zero, while the null for α is that the coefficient equals 1. The following denote statistical significance: *** 1% level, ** 5% level.