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Is employer-based health insurance a barrier to entrepreneurship?⋆

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ABSTRACT

The focus on employer-provided health insurance in the United States may restrict business creation. We address the limited research on the topic of "entrepreneurship lock" by using recent panel data from matched Current Population Surveys. We use difference-in-difference models to estimate the interaction between having a spouse with employer-based health insurance and potential demand for health care. We find evidence of a larger negative effect of health insurance demand on business creation for those without spousal coverage than for those with spousal coverage. We also take a new approach in the literature to examine the question of whether employer-based health insurance discourages business creation by exploiting the discontinuity created at age 65 through the qualification for Medicare. Using a novel procedure of identifying age in months from matched monthly CPS data, we compare the probability of business ownership among male workers in the months just before turning age 65 and in the months just after turning age 65. We find that business ownership rates increase from just under age 65 to just over age 65, whereas we find no change in business ownership rates from just before to just after for other ages 55–75. We also do not find evidence from the previous literature and additional estimates that other confounding factors such as retirement, partial retirement, social security and pension eligibility are responsible for the increase in business ownership in the month individuals turn 65. Our estimates provide some evidence that "entrepreneurship lock" exists, which raises concerns that the bundling of health insurance and employment may create an inefficient level of business creation.

1. Introduction

The predominant source of health insurance in the United States for working-age adults is employer-provided health insurance. Nearly two-thirds of adults under age 65 and three-quarters of all full-time workers have health insurance through employers (U.S. Census Bureau, 2007). A potential cost of this reliance on employer-provide health insurance is the non-portability of insurance across employers potentially resulting in "job lock." Workers may be reluctant to switch jobs when otherwise optimal because of the possible loss of coverage due to pre-existing condition exclusions, waiting periods on new jobs, loss of particular insurance plans, and disruption in the continuity of care with their healthcare providers.

Concerns about disruptions in health insurance coverage could also influence the decisions of individuals who are contemplating starting new businesses (Holtz-Eakin et al., 1996). Such individuals who are currently covered by employer-sponsored health insurance would eventually lose that coverage if they leave their job. Potential business owners could face high premiums in the individual health insurance market and the possibly prohibitive health costs of being uninsured. Furthermore, changes in health plans and providers may be disruptive and costly. New entrepreneurs may also be exposed to pre-existing condition limitations and waiting periods for coverage if they have a spell of uninsured unemployment between their employer-provided coverage and their new health insurance policy.[1] Unless they have alternative sources of health insurance coverage, such as through a spouse's employer,

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[1] The 1996 Health Insurance Portability and Accountability Act (HIPAA) mandates that pre-existing condition limitations and waiting periods cannot be imposed on individuals who had continuous prior health insurance coverage, but it does not apply to individuals who do not have continuous prior coverage.

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this health insurance conundrum may influence their decision to start a new business.

All of these factors suggest that the U.S. focus on employer-based health insurance may restrict the formation of new businesses and create the additional inefficiency of altering who becomes and who does not become an entrepreneur. Although concerns that health insurance costs are "killing new-business dreams" (Egerstrom, 2007; Keen, 2005) and that health insurance issues distort employment choices to the detriment of start-ups (Leonhardt, 2009; Baumol et al., 2007) have been voiced for several years, the issue has taken on a new salience with the passage of the Patient Protection and Affordable Care Act of 2010 (PPACA). In the debate leading up to the passage of PPACA, President Obama noted the concern:

This is something I hear about from entrepreneurs I meet – people who've got a good idea, and the expertise and determination to build it into a thriving business. But many can't take that leap because they can't afford to lose the health insurance they have at their current job.[2]

Under PPACA, states will create "exchanges" where individual consumers can purchase insurance, and insurers will not be able to apply pre-existing condition exclusions and price premiums based on health status.[3] Although these features of PPACA have the potential to weaken entrepreneurship lock, PPACA exempts existing health plans from most regulations potentially allowing disparities in the value of health coverage to persist for some time (Eibner et al., 2010).

Given these concerns, it is surprising that only a handful of studies have examined whether employer-provided health insurance limits business creation. The few studies in the literature find mixed results, with some estimating that health insurance reduces transitions into self-employed business ownership by as much as 25% and others finding no evidence that health insurance reduces business creation (Holtz-Eakin et al., 1996; Madrian and Lefgren, 1998; Bruce et al., 2000; Wellington, 2001; DeCicca, 2007). The lack of research on the topic contrasts sharply with a much larger literature that examines the effects of employer-provided health insurance on employer-to-employer mobility (see Gruber and Madrian, 2004, for a review).

In this paper, we address the lack of current research on the topic of "entrepreneurship lock" by providing a new study of whether the U.S. health insurance system impedes business creation. We use panel data created by matching consecutive years or months of the Current Population Survey (CPS) and two identification strategies to examine this question. First, following the identification strategy pursued in most analyses of job lock, we compare the probability of turnover of otherwise observationally equivalent employees who differ only in the value that they are likely to place on a current employer's health insurance policy. We estimate difference-in-difference models for the transition from wage-based employment to self-employed business ownership as a function of access to alternative health insurance and family health. Individuals with no alternative means of health insurance who obtain health insurance from their own jobs, and individuals who have poor family health should be less likely to become business owners. Our preferred set of models restricts the sample to workers with employer-provided health insurance, and compares transitions of workers without access to alternative health insurance and with poor family health to those who have access to alternative health insurance and have good family health. The CPS allows us to measure business creation of all types of businesses including incorporated, unincorporated, employer and non-employer businesses.

The second identification strategy exploits the abrupt change in health insurance coverage occurring at age 65 due to Medicare. The discontinuity in coverage suggests that a comparison of business ownership among individuals just below the age 65 cutoff to those just above the age 65 cutoff provides a test of the entrepreneurship lock hypothesis that is as close to a random experiment as possible. Although previous studies exploit the discontinuity in health insurance coverage created by Medicare (e.g. Card et al., 2008, 2009), the approach has not been previously taken to identify the effects of health insurance coverage on business creation. The lack of research on the topic may be due to the difficulty in finding a dataset with large enough sample sizes and a high-frequency measure of age. To address this problem, we use a novel procedure for identifying a person's age in months from matching monthly data from the CPS. To our knowledge, this is the first study using this procedure for identifying age in months from the CPS and the first study using the discontinuity created by Medicare to test the entrepreneurship lock hypothesis. The results from this new identification strategy and the difference-in-difference approach using recent data shed light on the question of whether employer-based health insurance restricts business creation in the United States.

2. Previous literature

The few studies that examine the relationship between business creation and an individual's health insurance coverage status find mixed results.[4] Holtz-Eakin et al. (1996) considered the effect of health insurance coverage status on transitions from employment to self-employed business ownership using the 1984–86 Survey of Income and Program Participation (SIPP) and the 1982–84 waves of the Panel Study of Income Dynamics (PSID). Their study used difference-in-difference models based on the notion that insured wage/salary workers who had families in poor health and workers who did not have access to spouse health insurance should be less likely to transition to self-employed business ownership. While their estimates were quantitatively large (a lack of health insurance portability stemming from employer-sponsored insurance reduced the probability of transition from employment to self-employment by 9–15% in the SIPP population), they were statistically insignificant. Therefore, the authors could not confirm that health insurance impeded transitions to business ownership. Madrian and Lefgren (1998) also examine the question using the SIPP and find that by using additional waves of SIPP data (1984–93), estimates of the effect of health insurance coverage status on transitions to self-employment attain statistical significance. In addition to using the difference-in-difference methodology used by Holtz-Eakin et al. (1996), they also use the passage of continuation of coverage mandates to identify the effect of health insurance coverage status on transitions to business ownership. Their estimates imply that a lack of health insurance portability accounts for a 25% reduction in business creation. In other work, Wellington (2001) uses a similar estimation methodology to analyze data from the 1993 Current Population Survey (CPS). The author estimates the impact of having health insurance through one's spouse on the likelihood of self-employed business ownership. Her estimates suggest that a guaranteed alternative source of health insurance would increase the probability of business ownership between 2.3 and 4.4 percentage points for husbands and 1.2 and 4.6 percentage points for wives.

[2] http://www.whitehouse.gov/the_press_office/Weekly-Address-President-Obama-Explains-How-Health-Insurance-Reform-Will-Strengthen-Americas-Small-Businesses/).

[3] http://www.allhealth.org/sourcebookcontent.asp?CHID=68.

[4] The literature on the effects of an individual's health status on entrepreneurship also find mixed results (see Parker, 2009).

Another potential source of variation in the health insurance market for business owners comes from the tax treatment of health insurance. The tax subsidy to health insurance for business owners, introduced at 25% in 1986 rose to 100% by 2003 in a number of discrete changes. Velamuri (2005) uses this variation and compares the female self-employment rate in 1984–85 to that in 1990–91 and finds that women with no spousal health insurance were substantially more likely (12–25%) to be self-employed when tax subsides were higher compared to women who had access to spouse health insurance. However, estimates based on transitions to business ownership were statistically insignificant. Selden (2009) also uses the variation provided by the increased tax subsidy to examine insurance rates for self-employed families in the Medical Expenditure Panel Survey. The results show substantial increases in private insurance for business owners and their spouses. Gumus and Regan (2008) present the raw percentages of workers transitioning into business ownership between 1995 and 2005 and find that the transition rate has been stable over time and does not show any evidence of increasing when tax credits were increased. Unlike Selden, they find no relationship between tax-deductibility and rates of health insurance coverage among business owners using the CPS. Studies using the variation provided by tax subsidies are likely to yield lower estimates than studies using other methodologies, because many small businesses have very low levels of sales and profits in the first few years of existence (U.S. Census Bureau, 1997), and thus are not eligible for or benefit only slightly from the tax credit.

DeCicca (2007) presents additional evidence on the effect of legislative changes on transitions to business ownership. The study focuses on the effect of New Jersey's 1993 Individual Health Coverage Plan that included an extensive set of reforms facilitating access to coverage that was not employer-linked. The results suggest that New Jersey's plan increased business ownership among New Jersey residents by about 15–25% – a large effect compared to the estimates obtained using the Tax Reform Act. On net, there appears to be little consensus in this literature on the existence or magnitude of the effect of health insurance on business creation.[5]

In this study, we use recent panel data created by matching consecutive years or months of the CPS to estimate the effect of health insurance coverage status on business creation. Most prior research on this topic uses data from the 1980s and early 1990s, however, many important changes have occurred in the health insurance and labor markets. In particular, health insurance costs have risen dramatically since the 1990s, particularly for small group and individual plans. The demographic composition of new business owners has also changed, with the near-elderly – a rapid-growing segment of the U.S. population and one that faces higher costs for individual health insurance plans – becoming more likely to consider business ownership (Zissimopoulos and Karoly, 2007). Finally, several state and federal initiatives have attempted to increase the portability of health insurance and lower the costs of insurance for business owners in recent years. To address these concerns a current examination of the role of health insurance in entrepreneurship is needed. We update previous research on the topic using more recent data and employ a new identification strategy to explore whether employer-based health insurance limits business creation.

[5] The literature on the effects of health insurance coverage on job mobility among wage/salary workers also finds mixed results. See Stroupe et al. (2001), Bradley et al. (2007), Sanz de Galdeano (2006) and Gilleskie and Lutz (2002) for a few recent examples, and Gruber and Madrian (2004) for a recent review.

3. Conceptual framework

Access to health insurance is a major concern among business owners. In a recent survey, health insurance costs were most frequently listed as the most critical problem faced by small businesses (National Federation of Independent Business, 2008). In a related survey, three-quarters of business owners listed cost as an important barrier to offering health insurance through their business and 78% rated the satisfaction with their premium costs as "low" (AWP, 2005). Furthermore, the burden of premium costs is disproportionately high on the smallest establishments – representing 5.7% of sales for solo practitioners compared to 2.8% for larger establishments (AWP, 2005).

Self-employed business owners who do not have alternative access to health insurance, such as through a spouse, may need to rely on the individual health insurance market. Premiums in the individual health insurance market can be very high. In 2009, the average annual premium for non-elderly single policies was $2985 and for family policies was $6328. These average premiums mask substantial variation across individuals. Notably, average premiums are substantially higher for older people ($5755 for single policies ages 60–64 and $9952 for family policies ages 60–64) (AHIP, 2009). It is also important to note that these averages are based on information from people who actually purchased policies in the individual market. Workers who leave an insured job have the option to continue group coverage through COBRA for up to 18 months by paying 102% of the premium. At $1111 a month for family coverage, COBRA is expensive and only a small fraction of those eligible to purchase COBRA coverage do so (FamiliesUSA, 2009).

In this section, we provide a formal conceptual framework to describe why the market for health insurance, as it currently exists in the U.S., might be a barrier to business creation. This framework provides a background for the empirical analysis that follows. This discussion is adapted from a model presented in Gruber and Madrian (2004).

We assume that all employer-sponsored group health insurance coverage is the same (health insurance is a homogenous good) and individuals either have it or they do not. Individuals have preferences over wage compensation (or the monetary return from self-employed business ownership) and employer-sponsored group health insurance.

A worker's utility can be described by $U_{ij} = U(W_{ij}, H_{ij})$, where U_{ij} is the utility of worker i at firm j. W_{ij} is the wage of worker i at firm j, and H_{ij} is a binary indicator of employer-sponsored health insurance coverage of worker i at firm j. Let ΔW_{ij} denote the compensating wage differential in firms offering health insurance reflecting the fact that if individuals value health insurance, they will accept a lower wage from an employer that offers health insurance. Firms face a cost, C_{ij}, of providing workers with health insurance. If self-employed individuals and firms could purchase insurance on a per-worker basis and this insurance was perfectly experience rated and wages were perfectly flexible, the compensating differential ΔW_{ij} would be equal to the cost of health insurance C_i. In this highly stylized model, health insurance would have no effect on the labor market equilibrium since self-employed individuals could purchase health insurance for the same cost as other employers. Workers pay the same compensating differential if they choose a job with insurance and as a result, they select a job or business ownership where they have the highest marginal product of labor. So, workers will switch from a job (j) with group employer-provided health insurance to self employment (s) with no group health insurance if $U(W_{ij} - \Delta W, 1) < U(W_{is}, 0)$. Self-employed business owners can then choose to purchase non-group coverage for a cost of C_i in the individual market. In this stylized model, wage earners who do not have employer-sponsored health insurance start a business based on a simple comparison of their marginal productivity in the two sectors, and therefore should be as likely to start a business than wage earners who have group health insurance.

This stylized model is not realistic in several ways. First, self-employed business owners face higher health insurance costs than large firms because of their inability to capitalize on economies of scale, higher administrative costs per person, and lower bargaining power with insurers.[6] A compensating wage differential could adjust for this factor (i.e. people would enter self employment only if the expected wage was higher in that sector), but it would still lead to distortions because some people have access to group health insurance (i.e. through a spouse's employer) while others do not. Second, employers cannot fully vary health insurance coverage and wages in accordance with each worker's insurance costs. Therefore, workers with high health costs may be paying far less than the true costs of their insurance under group insurance. This can lead to distortions because workers with high health costs will be less likely to leave to start businesses even if otherwise optimal. Finally, health insurance is not a homogenous good that can seamlessly be transferred from an employer to self-employment. Despite the HIPAA protections noted above, individuals may incur disruptions in their relationships with providers and changes in policy quality as a result of purchasing new insurance as a self-employed business owner. These aspects of the market for health insurance can lead to distortions in the level of business ownership, who starts a business, and the timing of starting a business over the life cycle.

Using the framework described earlier, even if an individual was less productive in job, j, with group health insurance than when self employed ($W_{ij} < W_{is}$), the individual may not choose self employment if $U(W_{ij} - \Delta W, 1) - U(W_{is}, 0) > 0$. In this case, the cost of forgoing group health insurance coverage outweighs the additional utility from higher wages under self-employment. Even though the individual can use the higher wages from self-employment to purchase individual insurance, this insurance is likely to have a substantially higher cost in the individual market, have lower quality, and/or pose a disruption in the continuity of care for the worker. We expect that wage earners for whom $U(W_{ij} - \Delta W, 1) - U(W_{is}, 0)$ is large will be less likely to move into self employment. This difference in utilities represents the value of group employer-provided health insurance relative to business ownership. This value will be lower for workers who have access to another source of health insurance (spouse, parent, government program) and it would also be lower for workers that would face relatively low insurance costs in the market for individual health insurance (young, healthy workers with few dependents). The end result is that some individuals may be dissuaded from starting businesses when it is otherwise optimal because of the link between health insurance and employment.

4. Data

We use data from the 1996 to 2006 Annual Demographic and Income Surveys (March) of the CPS. Each annual survey, conducted by the U.S. Census Bureau and the Bureau of Labor Statistics, is representative of the entire U.S. population and interviews approximately 50,000 households and more than 130,000 people. Although the CPS is primarily used as a cross-sectional dataset offering a point-in-time snapshot, it is becoming increasingly common to follow individuals for two consecutive years by linking surveys. Households in the CPS are interviewed each month over a four-month period. Eight months later they are re-interviewed in each month of a second four-month period. The rotation pattern of the CPS makes it possible to match information on individuals in March of one year who are in their first four-month rotation period to information from March of the following year, which represents their second four-month rotation period. This creates a one-year panel for up to half of all respondents in the first survey. To match the March CPS files from 1996 to 2006, we use the method discussed in Fairlie and London (2008). The supplemental samples to the 2001–2006 ADFs, which are generally not re-interviewed in the following March, are removed.

The main advantage of the matched CPS is the large sample size. The matched CPS sample that we use includes more than 160,000 observations for wage and salary workers in the first survey year. The sample includes 5100 transitions to self-employed business ownership, which is considerably larger than the other panel datasets such as SIPP and PSID. In their study of health insurance and entrepreneurship, Holtz-Eakin et al. (1996) report 700 transitions from the wage and salary sector to self-employment in their sample from SIPP and considerably less in the PSID.

Across, the 1996–2006 CPS surveys, we find that roughly 75% of CPS respondents in one survey can be identified in the subsequent year's survey. The main reason that match rates are less than 100% is because of the movement of individuals or households out of sampled dwelling units. The CPS does not follow individuals who move out of CPS sampled dwelling units in future months. Another problem is due to false positive matches. Although unique household and person identifiers are available in the CPS to match non-moving individuals over time, false matches occur because of miscoding. We use a procedure that compares the sex, race and age of the person in each March file to remove false matches. Any changes in coding are identified as false matches.[7] False match rates, however, are very low (roughly 3%) and do not vary substantially across years.

The loss of observations due to household movement raises concerns about the representativeness of the matched CPS sample. We investigate this issue further by conducting a comparison of mean values from the original cross-sectional CPS sample to means values from the matched CPS sample. As expected, we find that the matched sample has higher insurance, employment and marriage rates, and is more educated and older. The matched sample is also less likely to be a minority, live in the central city and receive public assistance. But, in all of these cases the differences are very small. For example, health insurance coverage rates are only 3% different and the matched sample is only one year older than the original sample (see Fairlie and London, 2008, for more details).

4.1. CPS health insurance measure

The CPS health insurance questions ask individuals to report all sources of health insurance coverage during the entire year prior to survey month.[8] However, comparisons of CPS estimates of health insurance coverage to other surveys that ask about insurance at the time of the survey reveal similar numbers. Estimates from the SIPP, MEPS and National Health Interview Survey (NHIS) indicate that roughly 40 million individuals were uninsured at the time of the survey in 1998 (CBO, 2003). CPS estimates for the number of individuals with no insurance for the entire year were also roughly 40 million in that year, suggesting that the CPS over-

6 http://www.rwjf.org/pr/synthesis/reports_and_briefs/pdf/no2_policybrief.pdf.

7 Age in the second survey year is allowed to be in the range from −1 to +3 from the first survey year.

8 The CPS asks separate questions about employer-provided (own and dependent), privately purchased, military, Medicaid, Medicare, Indian Health Service, and other sources of health insurance.

Table 1
Insurance type by business ownership or wage/salary work matched current population surveys (1996–2006).

	Uninsured (%)	Employer (%)	Employer dependent (%)	Individual (%)	Medicaid (%)	Medicare (%)	Other (%)	*N*
Men								
Self-employed business owners	20.8	32.7	21.1	21.0	1.0	1.8	1.5	16,480
Wage/salary workers	11.8	74.7	8.7	2.5	0.9	0.5	0.9	88,648
Other/not working	14.0	30.1	9.5	14.3	12.3	17.6	2.2	28,118
S.E. business owners (full-time)	20.1	33.7	21.5	21.1	0.9	1.3	1.4	14,905
Wage/salary (full-time)	10.8	77.1	8.4	2.1	0.6	0.3	0.8	81,560
Women								
Self-Employed business owners	19.0	19.0	34.6	22.1	1.6	1.5	2.2	7903
Wage/salary workers	10.5	62.4	20.3	3.2	1.8	0.6	1.3	85,286
Other/not working	14.5	13.7	32.4	14.3	12.0	11.1	2.0	57,974
S.E. business owners (full-time)	19.7	21.3	31.7	22.5	1.5	1.1	2.2	5782
Wage/salary (full-time)	9.4	69.0	16.7	2.4	1.1	0.3	1.1	69,725

Notes: (1) The sample includes individuals aged 25–64. (2) Self-employed business owners and wage/salary workers are defined as 20 or more weeks per year and 15 or more hours per week. Other/not working includes low hours workers and non-workers. (3) Full-time work is defined as 40 or more weeks per year and 30 or more hours per week. (4) Self-employed business ownership in the CPS captures all types of businesses including incorporated, unincorporated, employer and non-employer businesses.

Table 2
Insurance type by business ownership status matched current population surveys (1996–2006).

	Uninsured (%)	Employer (%)	Employer dependent (%)	Individual (%)	Medicaid (%)	Medicare (%)	Other (%)	*N*
Men								
New business owners	24.5	37.7	19.9	14.3	1.3	1.0	1.3	3377
Business owner in both years	18.6	31.8	22.8	22.7	0.9	1.6	1.5	11,742
Business ownership leavers	18.9	53.6	13.7	10.7	0.9	1.4	0.8	3460
Women								
New business owners	23.2	24.4	31.4	15.6	2.1	0.9	2.4	1803
Business owner in both years	17.4	19.2	35.3	23.7	1.3	1.4	1.7	4806
Business ownership leavers	16.7	35.7	28.2	14.8	1.6	1.3	1.8	1853

Notes: (1) The sample includes individuals aged 25–64 who work 20 or more weeks and 15 or more hours per week in both survey years. All observations with allocated class of worker, weeks or hours information are excluded from the sample. (2) New business owners are not self-employed in the first survey year, but are self-employed in the second survey year, and business ownership leavers are self-employed in the first survey year, but not the second survey year. (3) Self-employed business ownership in the CPS captures all types of businesses including incorporated, unincorporated, employer and non-employer businesses.

states the number of individuals who are uninsured for an entire year. Bhandari (2004) finds similar estimates of insurance coverage rates in the CPS and point-in-time estimates from the SIPP even within several demographic groups. Estimates from the SIPP and MEPS indicate the number of people who are uninsured for an entire year is between 21 and 31 million. Thus, CPS respondents may be underreporting health insurance coverage over the previous calendar year because of recall bias or because they simply report their current coverage (see Bennefield, 1996; Swartz, 1986; CBO, 2003; Bhandari, 2004, for further discussion). Even if the CPS estimates capture a point-in-time measure of health insurance coverage, the measure of health insurance status does not change from year to year and thus allows for an analysis of transitions in status.

5. Health insurance coverage and business ownership

Table 1 provides a descriptive profile of the variation in health insurance coverage by employment status.[9] We find that self-employed business owners are nearly twice as likely to be uninsured than wage/salary workers. Self-employed business owners in the CPS include owners of all types of businesses – incorporated, unincorporated, employer and non-employer firms. By defining ownership using the individual's main job activity, the CPS measure is more restrictive than the U.S. Census Bureau's measure of business ownership in the Survey of Business Owners (SBO), which includes consultants and side business owners (see Headd, 2005; Fairlie and Robb, 2008, for more discussion). Estimates from the CPS indicate that roughly 20% male and female business owners report no insurance compared to 11.8% of male wage/salary workers and 10.5% of female wage/salary workers. The uninsured rates for self-employed business owners are also higher than those for the other/not working population. Although this group includes the unemployed, not in the labor force and low hours workers, health insurance rates are 6.8 percentage points higher than rates for business owners for men and 4.5 percentage points higher for women.

Insured male business owners are most likely to get their coverage from employment (33%), followed by dependent employer coverage (21%) and individual coverage (21%). However, insured female business owners are most likely to get dependent employer coverage (35%), followed by individual coverage (22%) and coverage from own employment (19%). The distinction between individual coverage and own employer coverage for self-employed business owners is nebulous. Business owners may obtain health insurance only for themselves, but purchase it through their business, and report this coverage as employment-based insurance rather than individual insurance.

The lack of health insurance among full-time, full-year self-employed business owners is similarly high.[10] Slightly more than 20% of full-time, male business owners are uninsured and 19.7% of full-time, female business owners are uninsured. These rates of uninsurance are considerably higher than for full-time, wage/salary workers.

In Table 2, we use the two-year panel structure of our data to examine health insurance types and coverage in the second year

[9] Self-employment, hours worked, weeks worked and income are measured for the last calendar year to correspond to the health insurance variable.

[10] Full-time workers work 35 or more hours per week and 40 or more weeks a year.

Table 3
Wage/salary to business ownership transition rates by insurance type matched current population surveys (1996–2006).

	Wage/salary to business ownership entry rate (%)	*N*	W.S. to Bus. ownership entry rate (full-time) (%)	*N*
Men				
Total	4.0	83,061	3.7	74,505
Employer insurance	2.9	63,149	2.7	58,571
Employer dependent insurance	6.6	7437	6.7	6418
No insurance	6.5	8732	6.3	6893
Women				
Total	2.3	77,065	1.9	60,181
Employer insurance	1.5	49,511	1.3	42,847
Employer dependent insurance	3.2	15,366	2.9	9799
No insurance	3.7	7216	3.3	4852

Notes: (1) The sample includes individuals aged 25–64 who work 20 or more weeks and 15 or more hours per week in both survey years. All observations with allocated class of worker, weeks or hours information are excluded from the sample. (2) The full-time sample includes individuals aged 25–64 who work 40 or more weeks and 35 or more hours per week in both survey years. (3) Self-employed business ownership in the CPS captures all types of businesses including incorporated, unincorporated, employer and non-employer businesses.

for new business owners, business leavers, and business owners in both survey years. These estimates provide further evidence on the strong relationship between business ownership and not having health insurance. Individuals who are new business owners have very high rates of uninsurance – 24.5% for men and 23.2% for women – indicating that starting a business is strongly associated with the loss of health insurance. As reported in Table 1, both wage/salary workers and those not working had substantially lower rates of uninsurance.[11]

Although individuals who have owned a business for at least two consecutive years have higher rates of health insurance coverage than new business owners, coverage rates remain very low. Among men, 18.6% lack health insurance, and 17.4% of women are uninsured. Another interesting finding is that more than half of the male workers who leave business ownership move to jobs that have employer-provided health insurance. A large percentage of women leaving business ownership also move to jobs with employer-provided insurance. Overall, these results suggest that being uninsured is associated with movements to and from business ownership.

Four percent of all male wage/salary workers start a business each year (see Table 3). For those who have health insurance coverage from their employer, business creation rates are substantially lower at 2.9%. In contrast, 6.6% of workers who have health insurance coverage from a spouse start a business. Wage/salary workers who have no insurance coverage have a similarly high likelihood of starting a business. This result is not being driven by the unemployed or low-hours workers because only wage/salary workers with 20 or more weeks and 15 or more hours per week are included in the sample. Furthermore, when we condition on full-time, full-year work we find similar results. Business creation rates are substantially lower among wage/salary workers who have employer insurance than among wage/salary workers who have insurance coverage through a spouse or do not have insurance.

Although business entry rates are lower for women, similar patterns across health insurance coverage emerge. Business creation rates are much lower for female workers with employer insurance than for female workers with spousal coverage or no insurance. Conditioning on full-time work does not change this conclusion.

Of course, we cannot interpret these descriptive results as evidence that employer health insurance is an impediment to starting a business because employer-provided health insurance is correlated with job quality. Workers who have employer-provided health insurance may be less likely to start a business or switch to another job simply because they already have a job with a good compensation package. We attempt to address these concerns in the next section.

[11] Over half of the uninsured newly self-employed were insured before becoming self-employed, and for these workers the move to self-employment concurred with a loss of health insurance.

6. Estimating the effects of health insurance coverage status on business creation

We use two main estimation strategies to identify the effect of health insurance coverage status on business ownership. First, we construct difference-in-difference models of the transition to self-employed business ownership from wage-based employment as a function of access to alternative health insurance and family health. Individuals with no alternative means of health insurance who obtain health insurance from their own jobs, and individuals who have poor family health should be less likely to become business owners, all else equal. The second identification strategy takes advantage of the abrupt change in health insurance coverage occurring at age 65 due to Medicare. We explore whether the gain in health insurance at age 65 encourages individuals to become self-employed business owners by comparing rates of ownership among those just below age 65 with rates among those just above age 65.

6.1. Difference-in-difference estimates

The general approach taken here to identify the effect of health insurance coverage status on entrepreneurship is to compare the rate of business creation for an experimental group that potentially faces a disruption in health insurance coverage to the rate of business creation for a control group that does not face a disruption. In addition, we use the fact that groups with a high demand for their current health insurance policy should be less likely to leave their jobs to start a business. Previous studies taking this approach have used several different variables to proxy for high demand including number and health status of family members (Holtz-Eakin et al., 1996; Gruber and Madrian, 2004). We focus on a few of these measures that are available in the CPS and best capture potential demand for health insurance and care. The measures of potential health care demand that we include are the following: (i) having a family member in bad health, (ii) number of family members in bad health, and (iii) lacking an alternative source of health insurance coverage through a spouse's employer plan.[12] These measures of family bad health do not include the health status of the respondent.[13] Individuals who have a family member in poor health are likely to have a high demand for their current employer-provided health insurance policy since they may face high premiums in the individual health insurance market or a discontinuity in their treatment if they change insurance plans. Workers who have only a single source of employer-provided health insurance are likely to have a higher demand for this health insurance compared to workers who have access to an alternative source of health insurance from a spouse's employer-provided health insurance plan. Access to spouse's health insurance plan has been used in several previous studies of health insurance and business creation or job mobility (see Holtz-Eakin et al., 1996; Madrian and Lefgren, 1998; Madrian, 1994; Kapur, 1998; Wellington, 2001, for example).

While there is considerable flexibility in the choice of experimental and control groups in a difference-in-difference estimator, the comparability of the two groups is important to obtain a consistent estimator. The key assumption, which is likely to hold only if the groups are comparable, is that the effect of any exogenous influences is the same on the control and the experimental groups (Meyer, 1995). We use two main classifications of experimental and control groups. First, we define individuals who hold employer-provided health insurance as the experimental group and individuals who do not hold employer-provided health insurance as the control group. By definition, individuals who hold health insurance are more likely to be deterred from starting a business because of their current health insurance status than individuals who do not hold health insurance. Empirically, we estimate the following probit model:

$$\mathrm{prob}(y_i) = \Phi(\beta_0 + \beta_1 H_i + \beta_2 D_i + \beta_3 H_i D_i + \gamma' X_i) \tag{6.1}$$

where H_i denotes whether an individual holds employer-provided health insurance, D_i is potential health care demand, and X_i is a vector of demographic and job characteristics. The CPS allows us to include very detailed controls for the individual's job in the baseline year, family, individual demographics, residence, and survey year.[14] We estimate separate models for men and for women. The sample consists of wage and salaried workers in the baseline year (t). The dependent variable, y_i, equals 1 if the worker moves to self-employed business ownership in the following year ($t+1$). We estimate several versions of this model with the measures of potential health care demand discussed above. The coefficient on the interaction between health insurance and potential health care demand, β_3, captures the difference-in-difference estimate of "entrepreneurship lock."[15] A negative coefficient is consistent with the notion that current employer-provided health insurance is a disincentive to starting a business, and suggests that those individuals who would face a disruption in their health insurance and have a high demand for health care are relatively less likely to start businesses than individuals who have a low demand for health care. Note that we cannot simply interpret β_1 as the estimate of the effect of employer-provided health insurance on business creation because having own employer-provided health insurance may be correlated with high quality jobs and therefore this estimate would be biased.

Table 4A reports the results from estimating Eq. (6.1) for men using the full sample. Columns 1–3 present three different measures of high health care demand, no spouse health insurance, anyone in the family in bad health, and number of family members in bad health.[16] The estimates from the models in Table 4A show that whites and immigrants are more likely to start businesses. Workers with relatively more education, with higher family incomes and home-owners are also more likely to start businesses. In general, these results are consistent with findings from the previous literature and the notion that workers with more resources are the most likely to be able to start a business.[17]

The direct effect of own employer provided health insurance on the control group is large – workers who have such health insurance are between 2.5 and 3.9 percentage points less likely to start a business relative to a baseline transition rate of 4%. However, we cannot place much weight on the direct effect of health insurance since it could be contaminated by unobserved job quality, and so we rely on the interaction of the high demand variables with employer health insurance (e.g. β_3) to determine if insured individuals with a high demand for health care are relatively less likely to start businesses compared with individuals with a low demand for health care.

In column 1, the interaction of employer health insurance and no spouse health insurance is negative and statistically significant. The magnitude of the estimated effect is 2 percentage points which is quite large relative to a base business creation rate of 4% suggesting that the lack of spouse health insurance is a disincentive to starting a business for those who rely on their own employer policy. For the other measures of potential demand for health insurance in columns 2 and 3, the results are not as clear. The coefficients on the interactions between own employer health insurance and anyone with bad health and own employer health insurance and the number of family members with bad health are both negative, but statistically insignificant.

The results for women in Table 4B are somewhat similar. Employer provided health insurance has a large negative direct effect on business creation for the control group. It appears that higher wage women are also less likely to start businesses – the effects of wage and health insurance are similar for women, unlike for men. Similar to the results for men, the coefficient on the interaction between own employer health insurance and no spouse employer insurance is negative and statistically significant. The coefficient estimate is also large implying an effect of 1.75 percentage points. Using the alternative measures for potential demand, we do not find negative coefficients on the interaction terms.

A potential problem with this classification of experimental and control groups is that individuals who hold employer-provided health insurance differ from those who do not (Kapur, 1998). Insurance holders have higher wages, longer tenure, and more education than non-holders.[18] In additional specifications, we restrict the

[12] Bad or poor health is defined by individuals reporting that their health is "fairör "poorïnstead of "good,¨very good,ör "excellent.Špousal coverage is measured by using household, family and spouse identifiers for matching spouses, and information from each individual's employer health insurance coverage.

[13] The worker's own health is likely to have a strong effect on his own job choice, and is excluded for our main results. However, including own health provides similar results.

[14] The inclusion of survey year controls will capture any effect of changes in the tax treatment of health insurance over time. States also implemented insurance market reform; however almost all of these reforms were implemented before our data were collected.

[15] The marginal effects for interaction terms in a probit model may be biased (Ai and Norton, 2003). Results in the paper are very similar using a linear probability model. In addition, we have calculated predictions of the marginal effects and their distribution and found a similar pattern of results, although these are somewhat more cumbersome to report.

[16] We have also estimated the models with a measure of family health that includes the individual's own health. Results using this measure are quite similar to the results reported in the paper.

[17] See Parker (2009) and van Praag (2005) for recent reviews of the literature on the determinants of business ownership.

[18] In our data, insurance holders are paid $7 per hour more and are 15% more likely to have college degrees compared to non-holders. Among insurance holders, those who have spouse health insurance are almost identical to those who do not have it.

Table 4A
Probit regressions for probability of business creation for men matched current population survey (1996–2006).

Explanatory variables	(1)	(2)	(3)
Black	−0.0185	−0.0180	−0.0180
	(0.0033)	(0.0033)	(0.0033)
Latino	−0.0220	−0.0206	−0.0206
	(0.0034)	(0.0033)	(0.0033)
Asian	−0.0060	−0.0058	−0.0058
	(0.0041)	(0.0041)	(0.0041)
Immigrant	0.0119	0.0127	0.0127
	(0.0028)	(0.0027)	(0.0027)
Age	0.0019	0.0018	0.0018
	(0.0006)	(0.0006)	(0.0006)
Age squared	−0.0015	−0.0015	−0.0015
	(0.0007)	(0.0007)	(0.0007)
High school graduate	0.0073	0.0062	0.0062
	(0.0028)	(0.0028)	(0.0028)
Some college	0.0080	0.0066	0.0067
	(0.0029)	(0.0029)	(0.0029)
College graduate	0.0155	0.0139	0.0140
	(0.0031)	(0.0031)	(0.0031)
Graduate school	0.0215	0.0203	0.0204
	(0.0035)	(0.0035)	(0.0035)
Log wage	0.0039	0.0044	0.0044
	(0.0018)	(0.0018)	(0.0018)
Log family income	0.0053	0.0037	0.0037
	(0.0018)	(0.0018)	(0.0018)
Home ownership	0.0078	0.0070	0.0070
	(0.0020)	(0.0020)	(0.0020)
Own employer health insurance	−0.0253	−0.0392	−0.0392
	(0.0024)	(0.0016)	(0.0016)
No spouse employer health ins.	0.0211		
	(0.0026)		
Own employer HI × no spouse emp. HI	−0.0244		
	(0.0031)		
Anyone in family in bad health		−0.0045	
		(0.0038)	
Own employer HI × anyone bad health		−0.0003	
		(0.0050)	
Number in family in bad health			−0.0025
			(0.0029)
Own employer HI × number bad health			−0.0008
			(0.0040)
Mean of dependent variable	0.0398	0.0398	0.0398
Sample size	81,214	81,214	81,214

Notes: (1) The dependent variable equals 1 if the individual switches from wage and salary work in survey year t to self-employed business ownership in survey year $t+1$. (2) Marginal effects and their standard errors are reported. (3) All specifications include controls for other race, multiple race, marital status, children, spousal employment, interest income, dividend income, rental income, region, urbanicity, industry, and year of survey.

sample to individuals who hold employer-provided health insurance to improve the comparability of the experimental and control groups. We define the control group as individuals who have access to alternative health insurance from a spouse's employer. We do not require that the individual is covered by the spouse's plan, only that the spouse has own employer-provided health insurance, since individuals can usually obtain coverage from a spouse's employer even if they are not currently covered by the policy.[19] The experimental group is defined as individuals who do not have access to spousal employer-provided health insurance. Individuals who do not have access to an alternative plan should be more likely to be deterred from starting a business because of health insurance. Workers without spousal coverage face a potential disruption in health insurance coverage when moving from wage/salary work to business ownership, whereas workers with spousal coverage potentially do not face a potential disruption in health insurance. Individuals in these two groups are relatively similar across several dimensions such as wages, education, and tenure, suggesting that individuals with own and spousal employer-provided health insure form a more comparable control group for individuals with only employer-provided health insurance.[20]

We estimate the following probit model on the sample of individuals who hold employer-provided health insurance.

$$\text{prob}(y_i) = \Phi(\beta_0 + \beta_1 NS_i + \beta_2 D_i + \beta_3 NS_i D_i + \gamma' X_i), \qquad (6.2)$$

where NS_i denotes that an individual *does not* have a spouse who holds an employer-provided health insurance plan, and hence has a high demand for his own employer provided policy. The sample now only consists of wage and salaried workers in the baseline year (t) who hold employer-provided health insurance. The dependent variable equals 1 if the worker starts a business in the following year ($t+1$). We estimate this model with the remaining measures of potential health insurance demand. The coefficient on the interaction between no spousal health insurance and high health care demand, β_3, captures the difference-in-difference estimate of "entrepreneurship lock." As in Eq. (6.1), a negative coefficient sug-

[19] We do not have information on whether the individual was offered health insurance and turned it down.

[20] Individuals who have both employer-provided health insurance and access to spousal health insurance may still have a preference for their own employer policy, and as a result, prefer to stay in their current job. This would result in an underestimate of the effect of health insurance on business creation.

Table 4B
Probit regressions for probability of business creation for women matched current population survey (1996–2006).

Explanatory variables	(1)	(2)	(3)
Black	−0.0071	−0.0068	−0.0068
	(0.0024)	(0.0024)	(0.0024)
Latino	−0.0057	−0.0047	−0.0047
	(0.0026)	(0.0026)	(0.0026)
Asian	−0.0031	−0.0030	−0.0030
	(0.0032)	(0.0032)	(0.0032)
Immigrant	0.0063	0.0075	0.0075
	(0.0022)	(0.0022)	(0.0022)
Age	0.0006	0.0004	0.0004
	(0.0005)	(0.0005)	(0.0005)
Age squared	−0.0004	−0.0002	−0.0002
	(0.0006)	(0.0006)	(0.0006)
High school graduate	0.0076	0.0064	0.0064
	(0.0026)	(0.0026)	(0.0026)
Some college	0.0116	0.0102	0.0102
	(0.0027)	(0.0027)	(0.0027)
College graduate	0.0164	0.0152	0.0151
	(0.0029)	(0.0029)	(0.0029)
Graduate school	0.0219	0.0210	0.0210
	(0.0032)	(0.0032)	(0.0032)
Log wage	−0.0025	−0.0020	−0.0020
	(0.0012)	(0.0013)	(0.0013)
Log family income	0.0055	0.0035	0.0035
	(0.0012)	(0.0012)	(0.0012)
Home ownership	0.0100	0.0089	0.0089
	(0.0017)	(0.0017)	(0.0017)
Own employer health insurance	−0.0126	−0.0201	−0.0202
	(0.0017)	(0.0012)	(0.0012)
No spouse employer health ins.	0.0196		
	(0.0018)		
Own employer HI × no spouse emp. HI	−0.0175		
	(0.0023)		
Anyone in family in bad health		0.0011	
		(0.0026)	
Own employer HI × anyone bad health		0.0033	
		(0.0038)	
Number in family in bad health			0.0005
			(0.0021)
Own employer HI × number bad health			0.0036
			(0.0031)
Mean of dependent variable	0.0231	0.0231	0.0231
Sample size	75,317	75,317	75,317

Notes: (1) The dependent variable equals 1 if the individual switches from wage and salary work in survey year t to self-employed business ownership in survey year $t+1$. (2) Marginal effects and their standard errors are reported. (3) All specifications include controls for other race, multiple race, marital status, children, spousal employment, interest income, dividend income, rental income, region, urbanicity, industry, and year of survey.

gests that those individuals who would face a disruption in their health insurance and have a high demand for health care are relatively less likely to start a business than individuals with a low demand for health care.[21]

Table 5A reports estimates from Eq. (6.2) for men. We report the main effects and interactions between not having a spouse with employer health insurance and the two remaining health demand measures in Columns 1 and 2. The experimental group is defined as individuals who do not have spouses with employer health insurance and the control group is defined as individuals who have spouses with employer health insurance. The coefficient on the interaction between no spouse health insurance and anyone in bad health in the family is large, negative and statistically significant. The coefficient estimates on the number of family members in bad health is also negative and statistically significant. These estimates show that men with poor family health and no spouse health insurance are significantly less likely to give up their employer plan to start a business.

The results are similar for women (Table 5B). Female workers in families with poor health and do not have spouses with health insurance are less likely to start businesses. For both measures of poor family health the coefficients are large, negative and statistically significant.

6.2. Additional estimates

A concern with the specifications columns 1 and 2 of Tables 5A and 5B is that the experimental group includes both married and unmarried workers whereas the control group only includes married workers in dual-worker couples because this group has a spouse with employer health insurance. To further improve the comparability of the experimental and control groups we first limit the sample to married workers. We find that these results are similar to the ones reported in columns 1 and 2 of Tables 5A and 5B. To further increase comparability, we limit the sample to married couples with full-time, full-year working spouses (estimates are reported in columns 3 and 4 of Tables 5A and 5B). We find that the interaction between the health

[21] We also estimated models including a control for the interaction between marital status and health, to explore the possibility that the interaction term between spouse health insurance and health may be capturing the effect of being married. These models generated similar results.

Table 5A
Probit regressions for probability of business creation for men who have employer-provided health insurance matched current population survey (1996–2006).

Explanatory variables Sample restriction	(1) Only EPHI	(2) Only EPHI	(3) EPHI and spouse employed FT	(4) EPHI and spouse employed FT
No spouse employer health ins.	−0.0023	−0.0020	−0.0009	−0.0009
	(0.0018)	(0.0018)	(0.0021)	(0.0021)
Anyone in family in bad health	0.0061		0.0105	
	(0.0049)		(0.0056)	
No spouse employer HI × anyone bad health	−0.0134		−0.0183	
	(0.0057)		(0.0083)	
Number in family in bad health		0.0073		0.0103
		(0.0038)		(0.0044)
No spouse employer HI × number in bad health		−0.0141		−0.0163
		(0.0046)		(0.0068)
Mean of dependent variable	0.0290	0.0290	0.0306	0.0306
Sample size	62,060	62,060	30,596	30,596

Notes: (1) The sample includes only individuals with own employer provided health insurance in Specifications 1 and 2. (2) The dependent variable equals 1 if the individual switches from wage and salary work in survey year t to business ownership in survey year $t+1$. (3) Marginal effects and their standard errors are reported. (4) All specifications include controls for other race, multiple race, marital status, children, spousal employment, interest income, dividend income, rental income, region, urbanicity, industry and year of survey.

Table 5B
Probit regressions for probability of business creation for women who have employer-provided health insurance matched current population survey (1996–2006).

Explanatory variables Sample restriction	(1) Only EPHI	(2) Only EPHI	(3) EPHI and spouse employed FT	(4) EPHI and spouse employed FT
No spouse employer health ins.	0.0011	0.0010	0.0006	0.0006
	(0.0014)	(0.0014)	(0.0018)	(0.0018)
Anyone in family in bad health	0.0108		0.0105	
	(0.0029)		(0.0039)	
No spouse employer HI × anyone bad health	−0.0119		−0.0086	
	(0.0038)		(0.0062)	
Number in family in bad health		0.0089		0.0105
		(0.0024)		(0.0031)
No spouse employer HI × number in bad health		−0.0092		−0.0072
		(0.0032)		(0.0052)
Mean of dependent variable	0.0144	0.0144	0.0181	0.0181
Sample size	48,663	48,663	23,917	23,917

Notes: (1) The sample includes only individuals with own employer provided health insurance in Specifications 1 and 2. (2) The dependent variable equals 1 if the individual switches from wage and salary work in survey year t to self-employed business ownership in survey year $t+1$. (3) Marginal effects and their standard errors are reported. (4) All specifications include controls for other race, multiple race, marital status, children, spousal employment, interest income, dividend income, rental income, region, urbanicity, industry, and year of survey.

measures and spousal health insurance strengthens in magnitude slightly and continues to be statistically significant. However, for women, the interaction term becomes somewhat smaller in magnitude and statistically insignificant.

One limitation of the difference-in-difference model estimated on the sample of insured workers is the reliance on the assumption that spousal health insurance is exogenous to becoming a business owner (Royalty and Abraham, 2006). Although some spouses may actively look for jobs with health insurance when an individual starts a business it is unlikely that this drives the entire relationship. First, by conditioning on not owning a business in the first survey year we only allow spouses who are already employed at jobs with health insurance to contribute to the estimated effects of entrepreneurship lock. We do not capture the effect from spouses who find insured jobs simultaneously with or after the business creation decision. Second, there are a myriad of macroeconomic and labor market factors that affect the health insurance coverage of spouses that are unrelated to potential business creation.[22] We also do not use the direct effect of spousal insurance as a measure of the restriction on starting a business and instead use it to simply define the treatment and control groups, thus resulting in the potential for only a "second-order" bias whereby some of the treatment and control observations are misclassified. Finally, the regression discontinuity estimates presented in the next section, which do not rely on the assumption that spousal insurance is exogenous, do not contradict these results.

6.3. Regression discontinuity design estimates

In this section, we take a new approach to examining whether health insurance discourages business creation by exploiting the discontinuity created at age 65 through the qualification for Medicare. In the month that individuals turn 65, they automatically qualify for Medicare, providing universal access to health insurance coverage.[23] Card et al. (2008, 2009) show that health insurance coverage increases substantially at age 65. Attaining Medicare eligibility should immediately reduce the value an individual places on employer-sponsored health insurance. In particular, would-be-

[22] Chernew et al. (2005) find that rising health insurance costs account for changes in health insurance coverage. Other studies in this literature have found that industry shifts, increased reliance on part-time workers and crowd-out also explain part of the change.

[23] Individuals automatically qualify for Medicare Part A (hospital insurance) in the month they turn age 65 if they have 40 quarters of previously covered employment or have a qualifying spouse. Medicare Part B (insurance) can be purchased for a small monthly payment.

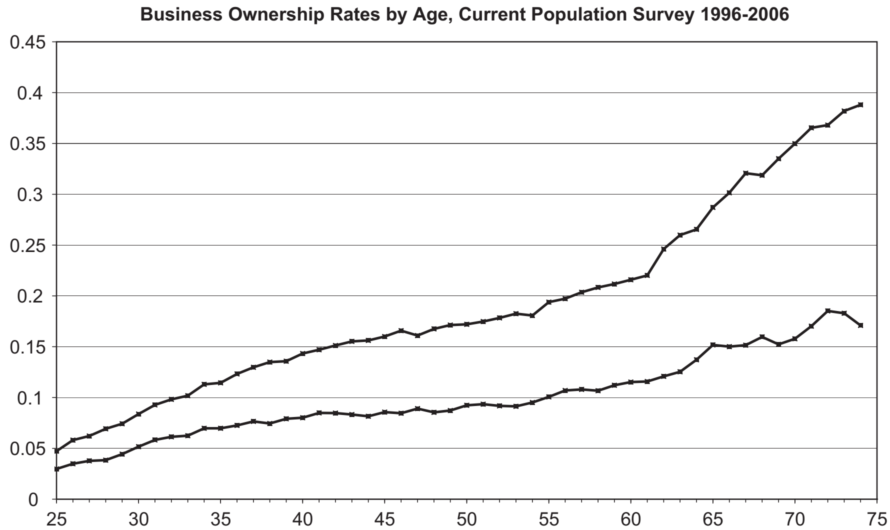

Fig. 1. Business Ownership Rates by Age, Current Population Survey 1996-2006.

entrepreneurs no longer have to be concerned about losing basic employer-sponsored health insurance coverage after that date. We can isolate the effects of the "Medicare notch" by comparing self-employed business ownership rates just before the age 65 birth month to just after the age 65 birth month. This approach addresses concerns over the potential influence of unobservables such as health insurance preferences and individual health status, on the results. The main criticism is that it provides only a local estimate of the effect of health insurance on self-employed business ownership for older workers. However, the recent debate over health care reform has included proposals to lower the eligibility age for Medicare.

To take this approach we use matched monthly data from the CPS.[24] By matching consecutive months of the CPS, we can identify the exact month in which the person's age in years changes. The CPS interviews households for 4 consecutive months, which allows us to identify up to two months before the birth month, the birth month, and 2 months after the birth month. We cannot, however, identify the birth month of individuals whose birth month does not fall in the four-month interview window. To our knowledge, this approach of using matched monthly CPS data has not been previously used to estimate regression discontinuity models. The approach is useful for identifying whether "entrepreneurship lock" exists. Few data sets contain a large enough sample size as well as information needed to identify exact birth month. The approach also has an advantage over many previous regression discontinuity studies that rely on age measured in years or quarters because we do not have to make potentially strong assumptions about the shape of the relationship between age and the outcome of interest. The effects of age on business ownership will be very small because we can zero in around the age 65 birth month.

The narrow focus on changes in the business ownership rate around the birth month is important for identification because business ownership rates increase substantially with age. Fig. 1 displays estimates for men between the ages of 25 and 74 in the workforce. Although there is a large increase in business ownership rates between age 64 and age 65, the rates increase steadily with age for older workers. The increase in the business ownership rate from 26.5% at age 64 to 28.7% at age 65 is the second largest increase. The largest increase in business ownership rates occurs between ages 61 and 62, but as discussed below we do not find evidence that the increase occurs in the month that the individual turns 62 coinciding with initial eligibility for social security benefits.[25] Although we do not focus on women, we also find a large increase in female business ownership rates at age 65. We do not focus on women in this analysis because the use of data from the 1996 to 2006 CPS implies that individuals who reach age 65 in the sample were born in the 1930s. There have been dramatic changes in labor force participation among women belonging to this cohort (Lichter and Costanzo, 1987; McEwen et al., 2005). In addition, this age group has a very low labor force participation rate. We find that only around 30% of women ages 55–75 in the sample are employed. We thus only examine business ownership for men in this section.

To focus the analysis around the month of the 65th birthday when individuals become eligible for Medicare we limit the sample to workers whose birth month falls in the four consecutive month interview period. We create three groups for the comparison of business ownership rates: the two months before a birth month (just under age 65), the month in which the age changes (possibly just over age 65), and the two months following an age change (just over age 65). The "possibly just over age 65" category is created because of the ambiguity over whether the individual's birthday is in the same calendar month as the survey month or if it falls in the calendar month after the survey month. The survey date is typically in the second week of the month. We also focus primarily on changes in the self-employed business ownership rate instead of transition rates. Unlike the previous section, we cannot model annual transition rates into business ownership because our empirical strategy requires us to compare consecutive months around the age 65 birth month. Although we can examine monthly transitions into business ownership, it is difficult to detect any statistically significant changes because the baseline monthly transition rate is extremely small at 0.004 and there are concerns

[24] A limitation of the basic monthly CPS data is that we have no information on health insurance coverage or health status. We thus cannot distinguish individuals by demand for health insurance or care.

[25] Initial eligibility for Social Security benefits may relax a liquidity constraint for some individuals wanting to start a business. While business ownership rates by age in years revealed an increase at age 62, when we carefully examined the breakpoint using monthly data both graphically and with regression analysis, there was no evidence of a statistically significant increase in business ownership at the 62nd birth month.

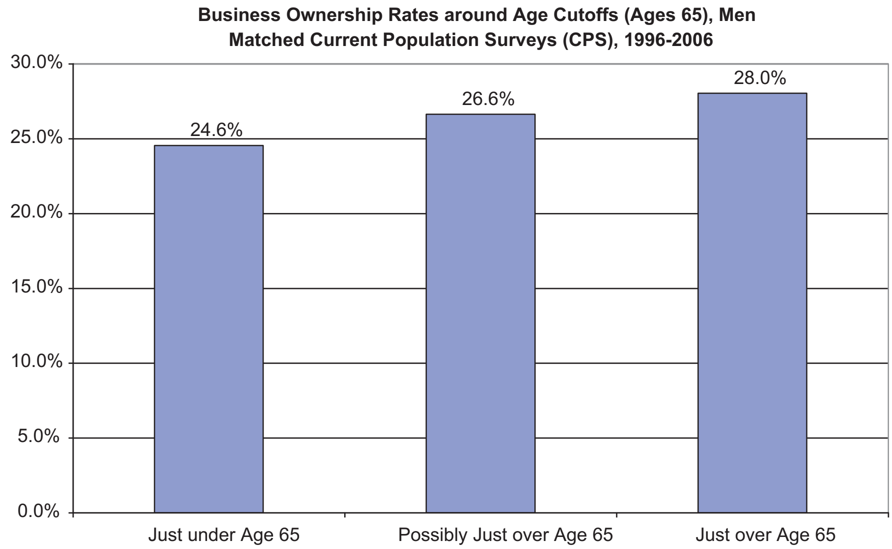

Fig. 2. Business Ownership Rates around Age Cutoffs (Age 65), Men Matched Current Population Surveys (CPS), 1996-2006.

over seasonal businesses.[26] As discussed below, we also estimate specifications for the transition rate and find roughly similar-sized point estimates.

Comparing business ownership rates around the 65th birth month indicates a clear break. Fig. 2 reports estimates of business ownership rates around the age 65 cutoff. Business ownership rates increase from 24.6% for those just under age 65 to 28.0% for those just over age 65. The difference is statistically significant. We also find that the business ownership rate increases from the just before age 65 category to the possibly age 65 category, which is consistent with an increase in rates in the month individuals turn age 65. Some individuals in the almost age 65 group will have turned age 65 by the survey date.

The increase in business ownership rates from two months before turning 65 to two months after turning 65 does not appear to be due to the slight increase in age. Fig. 3 reports estimates of business ownership rates from just before to just after changes in other ages 55–75. As expected, because age is only increasing slightly from just before to just after the birth month the business ownership rates are essentially the same around the birth month cutoff. For these age changes, there is no change in eligibility for health insurance. Additionally, when we examine changes in business ownership rates from just before to just after for each age in years we only find two ages with statistically significant changes other than the age 65 break.[27] Neither of these, however, was as large as the age 65 break. Note that given the 20 tests conducted, we expect to find a couple of statistically significant differences simply by chance and finding only two is reasonable given a 5% level of significance.

The discontinuity at age 65 can also be seen from a plot of business ownership rates before and after the cutoff using our sample of workers with birth months falling in the four-month window of the CPS (see Fig. 4). Separate linear predictions on either side of the discontinuity indicate a break in business ownership rates at age 65. The predicted difference in rates is similar in magnitude to the actual break in rates from just before to just after turning age 65.

To further investigate the discontinuity at age 65, we estimate regressions in which we control for demographic and job characteristics. We start with a simple regression for the probability of business ownership among workers who are just under or just over age 65. We model the probability of business ownership as:

$$\text{prob}(y_{it}) = \Phi(\alpha + \lambda_t + \delta_1 D_i^{65a} + \delta_2 D_i^{65o} + \beta' X_i), \tag{6.3}$$

where λ_t are year fixed effects, $D_i{}^{65a}$ is a dummy for possibly being age 65, $D_i{}^{65o}$ is a dummy for being just over age 65, and X_i is a vector of demographic and job controls. The omitted group is just under age 65. Identification of the Medicare effect, δ_2, is being driven entirely by comparing the just over age 65 group to the just under age 65 group. Because the potential effect of the slight increase in age on business ownership is likely to be small, the results from this specification are likely to be very robust to changes in the sample range and controlling for age.

Table 6 reports estimates from several regressions of Eq. (6.3). The first specification includes only the age 65 cutoff dummy variables. The omitted category is just under age 65. The coefficient on the just over age 65 variable is positive and statistically significant. The coefficient on the possibly age 65 variable is smaller, but statistically insignificant. We also include controls for race, nativity, education, marital status, region, urban status, industry and year in the remaining specifications. The coefficient estimate on the just over age 65 variable remains very similar attesting to the strength of the research design. The addition of the covariates has little effect on the estimated relationship between being just over the age 65 cutoff and business ownership.

We also estimate regression discontinuity models expanding the sample and controlling explicitly for age. In this case, the probability of business ownership is:

$$\text{prob}(y_{it}) = \Phi(\alpha + \lambda_t + g(a_i) + \delta_1 D_i^{65a} + \delta_2 D_i^{65o} + \delta_3 D_i^{a} + \delta_4 D_i^{o} + \beta' X_i), \tag{6.4}$$

where $g(a)$ is a function of age in months, $D_i{}^{a}$ is the "possibly age" dummy for all age groups, and $D_i{}^{o}$ is the just over dummy for all age

[26] The difficulty of identifying statistically significant changes in very small proportions has been noted previously (see Cohen, 1988 for example). A simple comparison of t-statistics for the test of 10% changes in the monthly transition rate into self-employment and the self-employment rate, which is 0.23 reveals that a roughly 10 times larger sample size is needed to find a statistically significant change in the monthly transition rate.

[27] The other breaks occur at age 59 and 61. We do not have a theoretical explanation for why business ownership rates would increase in these birth months and they may be due to chance.

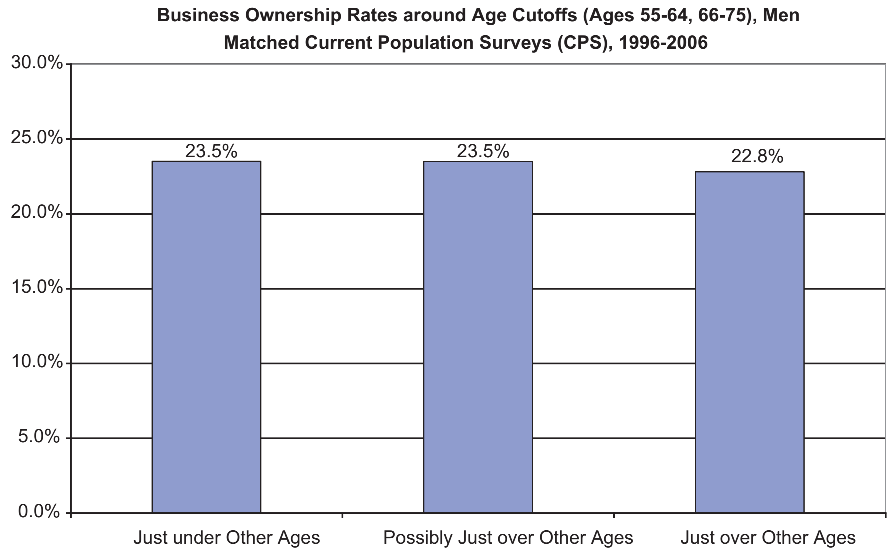

Fig. 3. Business Ownership Rates around Age Cutoffs (Ages 55-64, 66-75), Men Matched Current Population Surveys (CPS), 1996-2006.

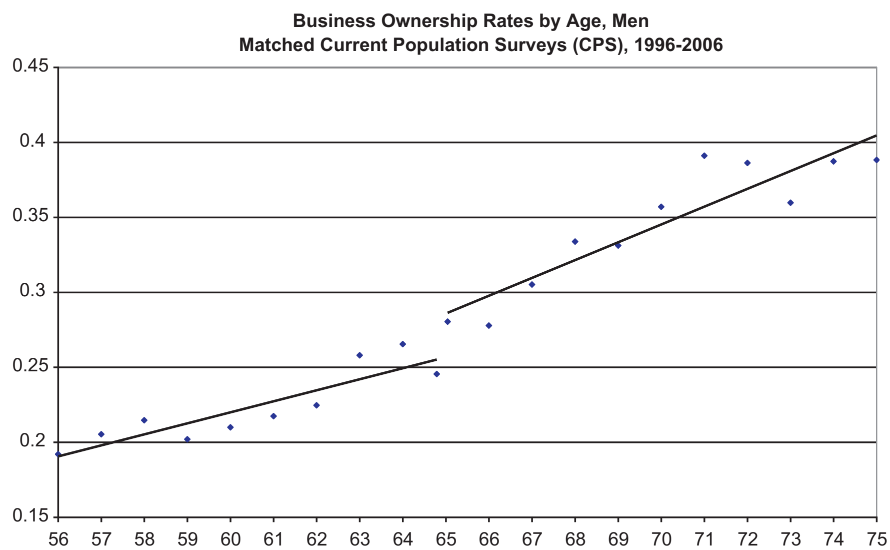

Fig. 4. Business Ownership Rates by Age, Men Matched Current Population Surveys (CPS), 1996-2006.

Table 6
Probit regressions for probability of business ownership, men around age 65 matched current population surveys (1996–2006).

Explanatory variables	(1)	(2)	(3)	(4)
Possibly just over age 65	0.02112	0.01893	0.01998	0.01965
	(0.01668)	(0.01623)	(0.01591)	(0.01523)
Just over age 65	0.03493	0.03280	0.03286	0.03029
	(0.01732)	(0.01690)	(0.01651)	(0.01587)
Year fixed effects	No	No	No	Yes
Demographic controls	No	Yes	Yes	Yes
Industry controls	No	No	Yes	Yes
Mean of dependent variable	0.26329	0.26329	0.26329	0.26329
Sample size	4015	4015	4015	4015

Notes: (1) The sample consists of workers around age 65 with 15 or more hours worked per week. (2) Demographic controls include race, nativity, education, marital status, region, and urban status.

groups.[28] We include these dummy variables for all age groups to rule out of the effect of the slight increase in age and the possibility of a birthday month effect on business creation. Similar to Eq. (6.3) the omitted group is just under age 65. By including the just over dummy for all ages we capture the effects of the small change in age associated with being just over to just under a specific age. We also estimate regressions with two general forms for $g(a)$. First, we estimate a standard quadratic form for age. Second, we estimate

[28] We adjust the standard errors for clustering by age group.

Table 7
Probit regressions for probability of business ownership, men matched current population surveys (1996–2006).

Explanatory variables	(1)	(2)	(3)	(4)
Possibly just over age cutoff	−0.00017	−0.00148	−0.00039	0.00000
	(0.00298)	(0.00297)	(0.00332)	(0.00310)
Just over age cutoff	−0.00710	−0.00219	−0.00047	−0.00110
	(0.00828)	(0.00525)	(0.00340)	(0.00317)
Possibly just over age 65	0.02002	0.01982	0.01998	0.01637
	(0.00298)	(0.00297)	(0.01591)	(0.01484)
Just over age 65	0.03992	0.03222	0.03286	0.03122
	(0.00828)	(0.00527)	(0.01651)	(0.01537)
Age quadratic	No	Yes	No	No
Age in year dummies	No	No	Yes	Yes
Year fixed effects	No	No	No	Yes
Demographic controls	No	No	No	Yes
Industry controls	No	No	No	Yes
Mean of dependent variable	0.23382	0.23382	0.23382	0.23382
Sample size	102,027	102,027	102,027	102,027

Notes: (1) The sample consists of workers aged 55–75 with 15 or more hours worked per week. (2) Standard errors are adjusted for clustering by age in years. (3) Demographic controls include race, nativity, education, marital status, region, and urban status.

a very flexible form for $g(a)$ that includes age in year fixed effects instead of a smooth function. This model is more flexible than most regression discontinuity models because it allows the pre and post age 65 levels to vary fully by age in years.

In Table 7, we report estimates for Eq. (6.4). We now include observations for all workers ages 55–75. The first specification includes only the age cutoff dummy variables. The omitted category is just under age 65. The coefficient on the just over age 65 variable is positive and statistically significant. Although there is a strong positive association between business ownership and age, the results are not being driven by the small increase in age from the just before period to the just after period. We are implicitly controlling for this increase in age by including dummy variables for possibly at the age cutoff and just over the age cutoff for all ages. As expected, these coefficients are very small suggesting that the small change in age between these two periods for ages other than 65 when individuals qualify for Medicare does not have an effect on business ownership. Nevertheless, we estimate additional specifications with further controls for age and other variables to check the robustness of the results. In Specification 2, we include a quadratic function for age in months. The coefficient estimate on just over age 65 remains large, positive and statistically significant.

In Specification 3, we replace the quadratic function for age in months with a specification that includes dummies for each age in years. Allowing for this more flexible form for the age–business ownership relationship, the estimates remain similar. We find a 0.033 higher probability of owing a business each month if the person is just over age 65 than if the individual is just under age 65. Finally, we also include controls for year, race, nativity, education, marital status, region, urban status, and industry in Specification 4. The coefficient estimate on the just over age 65 variable remains very similar providing further evidence on the credibility of the regression discontinuity design. The addition of the covariates has little effect on the estimated relationship between being just over the age 65 cutoff and business ownership. For this specification, the coefficient estimate implies a 0.031 higher probability of owing a business each month if the person is just over age 65 than if the individual is just under age 65.[29] This increase represents 13% of the mean probability of business ownership.

[29] Including dummy variables for other age breaks we only find significant coefficients for age 59 and 61, but both have smaller coefficients than the age 65 break. This is similar to the univariate results noted above.

6.4. Additional estimates and potentially confounding factors

We next investigate whether the estimates reported above are sensitive to sample and definitional changes, and whether there exist other confounding factors that lead to changes in work behavior in the month that individuals turn 65. We first narrow the age range to 60–70 year olds. Specification 1 of Table 8 reports estimates using a sample with this age range. The coefficient estimate on the just over age 65 variable remains large, positive and statistically significant. We try additional age ranges and find robust results.

One possible concern is that the change in the probability of business ownership observed in these results may be due to changes in composition of the labor force or due to transitions other than the move from wage/salaried work to full-time business ownership. For instance, wage/salaried workers may be moving to part-time self-employed business ownership at age 65 as part of their transition to retirement. To determine if this is driving our results, we restrict the sample to include only full-time workers (defined as working 30 or more hours per week). This restriction rules out the possibility that movement to part-time business ownership at age 65 is driving the results. As reported in Specification 2, the coefficient estimate is similar to the original one. Another possibility is that the stock of workers falls at age 65, and therefore the number of business owners as a share of the total workforce appears to be increasing even though the number of self-employed workers remains constant. To address this concern, we expand the sample to include individuals who are not working 15 or more hours per week. We now include all individuals aged 55–75 even if they are not in the labor force to ensure that the denominator is not affected by the size and composition of the labor force. The probability of business ownership for this sample is lower (11.0%) because of the inclusion of non-workers. Specification 3 of Table 8 reports estimates using this sample. We find a higher rate of business ownership associated with being just over the age 65 break. The point estimate implies that the business ownership rate is 0.013 higher, which represents 12% of the sample mean. The relative magnitude of the coefficient is similar to the coefficient estimate using the main sample of workers. Thus, the results do not appear sensitive to the treatment of non-employment and low hours work. We also estimated a regression in which hours worked was the dependent variable and found no change in hours worked around the age 65 cutoff. The coefficient estimate on the just over age 65 variable was very small and statistically insignificant.

To further investigate whether individuals are retiring or dropping out of the work force in the month they turned age 65, we

Table 8
Probit regressions for probability of business ownership, men, additional estimates matched current population surveys (1996–2006).

Explanatory variables Dependent variable	(1) SE rate	(2) SE rate	(3) SE rate	(4) SE transition
Possibly just over age cutoff	−0.00117	−0.00088	0.00002	−0.00010
	(0.00507)	(0.00324)	(0.00151)	(0.00037)
Just over age cutoff	−0.00347	0.00122	−0.00043	−0.00011
	(0.00525)	(0.00331)	(0.00155)	(0.00038)
Possibly just over age 65	0.01921	0.01309	0.00727	0.00012
	(0.01622)	(0.01587)	(0.00706)	(0.00167)
Just over age 65	0.03649	0.03525	0.01344	0.00116
	(0.01681)	(0.01640)	(0.00726)	(0.00164)
Age in year dummies	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Demographic controls	Yes	Yes	Yes	Yes
Industry controls	Yes	Yes	Yes	Yes
Mean of dependent variable	0.26019	0.22404	0.10992	0.00396
Sample size	43,797	91,083	215,052	183,871

Notes: (1) The sample consists of workers aged 60–70 in Specification 1, full-time workers aged 55–75 in Specification 2, and all individuals aged 55–75 in Specification 3, and non-business owners aged 55–75 in Specification 4. (2) Demographic controls include race, nativity, education, marital status, region, and urban status.

estimate a model in which employment is the dependent variable. We find a small and statistically insignificant coefficient estimate on the just over age 65 variable. This finding is consistent with estimates reported in Card et al. (2008, 2009) and von Wachter (2009).[30] The lack of empirical evidence that turning 65 affects retirement is somewhat puzzling given the positive effects on business ownership. One possibility is that an increasing share of the near-elderly who have a preference for retiring before 65 attain early Medicare eligibility via Disability Insurance and are retired before their 65th birthday.[31] Another possibility is that the effect of health insurance at the time of the 65th birth month is small relative to other factors affecting the retirement decision such as lost income, employment contracts, and family and health issues.[32]

A major concern with the regression discontinuity estimates is that there might exist other confounding factors that lead to shifts in employment behavior at age 65 such as eligibility for Social Security or pensions. Zissimopoulos and Karoly (2008) find that among individuals over age 50, those who had experienced a pension cash-out were more likely to transition from wage and salary work to self-employed business ownership. Social Security eligibility does not appear to generate shifts in employment behavior precisely at age 65. The minimum retirement age for full Social Security benefits was 65 for individuals born in 1937 or earlier (i.e., those reaching eligibility before 2003) and is gradually increasing for later birth cohorts to age 67 for those born after 1959. The earliest age of eligibility for Social Security benefits is 62; benefits received by individuals at that point are reduced (in an actuarial neutral way) relative to what would be received if one were to retire at the full retirement age. Data reveal that individuals are far more likely to begin claiming benefits at age 62 than at age 65. A majority of Americans (59% of women and 56% of men) receiving Social Security benefits for the first time in 2004 were age 62. A smaller fraction of those claiming benefits (17% of women and 23% of men) were age 65 (Munnell and Sass, 2007).

Similarly, age 65 does not appear to be a primary focal point for the accrual or availability of pension wealth. Under defined contribution retirement plans, pension wealth accrual does not vary substantially by age; pension wealth continues to increase as long as a person works. The critical age for individuals covered under defined contribution plans is 59.5 because at that age individuals can begin withdrawing from a 401(k) without penalty (Friedberg and Webb, 2003). Under defined benefit plans, pension wealth accrual peaks at the age of early retirement eligibility, which is well before age 65. Pension wealth may continue to increase up to age 65 (Friedberg and Webb, 2003; Poterba et al., 2001). In both the case of Social Security and pensions the evidence provided in previous studies does not indicate a major change in take-up at age 65.

The final robustness check involves focusing on transitions from non-business ownership to business ownership. As noted above, we cannot examine annual transition rates into self-employed business ownership similar to the analysis using the matched March CPS files because our empirical strategy requires us to compare consecutive months. Instead, we can only examine monthly transition rates which have a very low probability (sample average = 0.004). Nevertheless, we estimate Eq. (6.4) using the monthly business entry as the dependent variable as a robustness check. Specification 4 of Table 8 reports estimates. Similar to the previous results, we find a positive coefficient estimate on the just over age 65 variable. The coefficient, however, is not statistically significant. The point estimate implies that the business entry rate is 0.001 higher, which represents 29% of the sample mean. The magnitude of this coefficient estimate relative to the mean is larger, but roughly consistent with the finding for the just over age 65 variable in the business ownership rate specifications.[33]

7. Conclusions

A major concern with the U.S. focus on employer provided health insurance is that it might restrict business starts. The potential loss or disruption in health insurance coverage due to pre-existing condition limitations, waiting periods for coverage, changes in health plans and providers, high premiums in the individual health insurance market, and risk of high health costs while uninsured may dissuade many employees from starting a business when it would otherwise be optimal. Given these concerns it is surprising that only a handful of studies have examined whether employer-provided health insurance limits entrepreneurship, with

[30] Card et al. (2008, 2009) also do not find evidence of changes in marriage, family income and household moves at age 65.

[31] Autor and Duggan (2006) have shown increasing disability rolls and a looser definition of qualifying disabilities over time.

[32] Retiring workers face loss of income and health insurance. Workers moving to self-employment are likely to maintain some income, but lose health insurance. Thus, obtaining health insurance through Medicare may be a relatively bigger factor for workers moving to self-employment compared to retiring workers.

[33] We also estimated a regression for the monthly probability of exiting from self-employment and found a very small and statistically insignificant coefficient estimate on the just over age 65 variable.

the few studies in this literature finding mixed results. We address the limited research on the topic of "entrepreneurship lock" by providing a new study using panel data created by matching consecutive years or months of the CPS and two main identification strategies – difference-in-difference and regression discontinuity models. A first pass at the data reveals that self-employed business owners are much less likely to have health insurance than are wage/salary workers and even our sample of unemployed and part-time workers. Estimates from our two-year panel data from matching consecutive March CPS files also indicate that new business owners have especially low rates of health insurance coverage. We also find that business creation rates are substantially lower among wage/salary workers who have employer insurance than among wage/salary workers who have insurance coverage through a spouse or do not have insurance.

To address concerns that workers who have employer-provided health insurance may be less likely to start businesses because they already have a job with a good compensation package and high job quality, we first estimate difference-in-difference models based on the approach taken in the previous literature (e.g. Holtz-Eakin et al., 1996; Madrian, 1994). Identification of "entrepreneurship lock" arises from the interaction between having employer-provided health insurance and potential demand for health care. Using this first approach, we find some evidence that employer-based health insurance limits business creation, especially for men, but the evidence is not consistent across different measures of potential demand for health care. To improve the comparability of the experimental and control groups, we limit the sample to only individuals who have employer-based health insurance. Identification then comes from the interaction between having a spouse with employer-based health insurance and potential demand for health care. For men, we find consistent evidence of a larger negative effect of health insurance demand on the business creation probability for those without spousal coverage than for those with spousal coverage. Several robustness checks that further refine the comparability between experimental and controls groups provide similar results. Our estimates suggest that "entrepreneurship lock" for men is just over 1 percentage point relative to an annual base business creation rate of 3%. We also find evidence of entrepreneurship lock for women, however, the coefficients are not precisely estimated in a couple of specifications.

We also take a new approach in the literature to examining the question of whether employer-based health insurance discourages business creation by examining the discontinuity created at age 65 through the qualification for Medicare. Using a novel procedure of identifying age in months from matched monthly CPS data, we compare the probability of business ownership among male workers in the months just before turning age 65 and in the months just after turning age 65. Business ownership rates increase from 24.6% for those just under age 65 to 28.0% for those just over age 65, whereas we find no change in business ownership rates from just before to just after for the remaining ages in our sample of workers ages 55–75. We estimate several regression discontinuity models to confirm these results. As expected because of the small change in actual age and the orthogonality of included controls, we find a similarly large and statistically significant increase in business ownership rates in the age 65 birth month when the worker qualifies for Medicare. These results are not sensitive to several alternative samples, dependent variables, and age functions, and we do not find evidence from previous studies and additional specifications that other factors such as retirement, partial retirement, social security and pension eligibility are responsible for the increase in business ownership rates in the month the individual turns 65 and qualifies for Medicare.

Estimates from the difference-in-difference and regression discontinuity models both provide evidence that the U.S. emphasis on employer-provided health insurance may be limiting the creation of small businesses and influencing the decisions of workers regarding whether and when to start businesses. Our findings are consistent with the argument that relatively low rates of business ownership in the United States may be due to less comprehensive health insurance coverage than in other wealthy countries (Schmitt and Lane, 2009) and that expanding health insurance coverage will encourage business creation (Gruber, 2009). The recently enacted PPACA stipulates that individuals will be able to purchase insurance from insurance exchanges. Insurers will not be allowed to have pre-existing condition exclusions or premiums priced on the basis of health status. These features of PPACA may encourage business creation by providing potential entrepreneurs with a health insurance option should they leave their current employment. However, it remains unclear what the relative value of that option will be. Because PPACA exempts existing health plans from certain regulations, a disparity between the value of health coverage through the exchanges and the value of coverage through some existing employer plans is likely to persist for some time (Eibner et al., 2010). Moreover, the value of insurance provided through the exchanges will be influenced by the way in which states choose to structure them. PPACA will be phased in over the next few years with the availability of a high risk pool for the purchase of insurance in 2010 and ultimately the option of insurance exchanges in 2014. Investigating the impact of these changes on the health insurance market and entrepreneurship is an important area for future research.

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