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The Effects of Market Structure and Payment Rate on the Entry of Private Health Plans into the Medicare Market

Private insurance firms participating in Medicare can offer up to three principal plan types: coordinated care plans (CCPs), prescription drug plans (PDPs), and private fee-for-service (PFFS) plans. Firms can make entry and marketing decisions separately across plan types and geographic regions. In this study, we estimate firm-level models of Medicare private plan entry using data from the years 2007 to 2009. Our models include a measure of market structure and separately identify CCP, PDP, and PFFS entry. We find evidence that entry barriers associated with CCP market concentration affect all three product types. We also find evidence of cross-product competition and common cost or demand factors that make entry with certain product combinations more likely. We predict that the market presence of CCPs and PFFS plans will decrease and that of PDPs will increase in response to payment reductions included in the new health reform law.

Key questions in applied industrial organization pertain to the role of market structure in firms' entry decisions. A different set of inferences may be drawn about a market if concentration encourages entry rather than hinders it. On one hand, high market concentration suggests the existence of markups and a profit opportunity for potential entrants. On the other hand, incumbent firms may possess high market concentration due to barriers that preclude profitable entry of additional firms (Amel and Liang 1997). Thus, estimating the effect of market structure on entry provides a useful test of the presence of entry barriers. To our knowledge, no empirical analysis has been conducted that relates market structure to entry into the current, multiproduct Medicare private plan market.

In this study, we estimate reduced-form, firm-level entry models for the three main Medicare private plan types: coordinated care plans (CCPs) (largely health maintenance organizations [HMOs] and preferred provider organizations [PPOs]), private fee-for-service (PFFS) plans, and stand-alone prescription drug plans (PDPs). In doing so, we investigate the relation between market structure

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and entry for the multiproduct Medicare private plan market, controlling for the endogeneity of market structure with an instrumental variables (IV) approach. We focus on a select subset of plans that have exhibited the interest and ability to offer all three plan types. We find that higher CCP market concentration decreases the entry probability of that plan type, consistent with the presence of barriers to entry for the CCP submarket. We also find that CCP market concentration decreases the entry probability of the other two plan types, PFFS plans and PDPs, which is consistent with joint decision making across products within firms, possibly due to economies of scope.

We use our models to simulate changes in entry that would be induced by Medicare Advantage plan payment provisions in the new health reform law—the Patient Protection and Affordable Care Act (Pub. Law 111-148), as modified by the Health Care and Education Reconciliation Act of 2010 (Pub. Law 111-152) (hereafter jointly referred to as the ACA). Our simulations predict substantial market exit for the two main Medicare Advantage plan types that offer comprehensive benefits—CCPs and PFFS plans—and modest plan entry for PDPs.

These conclusions cannot be drawn from other Medicare plan entry models in the literature. To our knowledge, no recent paper has studied Medicare plan entry from the firm's perspective and none has estimated entry models across multiple product types. Our approach addresses these shortcomings. Because PDP entry decisions involve a different response to market characteristics in general and Medicare Advantage payment rates in particular, distinguishing PDPs from other plan types is a significant innovation in modeling entry of private plans into Medicare.[1] The same can be said for distinguishing PFFS plans and CCPs, although these plan types have more in common. Our models capture the differential effects of market characteristics on firms' entry decisions for these products by modeling them as distinct.

PDPs are a relatively new Medicare plan type, created within Medicare Part D[2] by the passage of the Medicare Prescription Drug, Improvement, and Modernization Act of 2003 (MMA; Pub. Law 108–173). In 2006, PDPs joined CCPs and PFFS plans, which had been available under Medicare Part C.[3] These older Medicare Advantage plan types offer comprehensive coverage of all Medicare services for a predetermined monthly payment from Medicare that is adjusted for the enrollee's expected spending. The main difference is that CCPs use a network of contracted providers whereas PFFS plans do not.[4]

The MMA also increased payments to Medicare Advantage plans, contributing to an increase in the number of plans offered and the number of beneficiaries covered (Pizer, Frakt, and Feldman 2009). PFFS plans experienced the most rapid increase in popularity following passage of the MMA (Frakt, Pizer, and Feldman 2009). Commensurate increases in the Medicare budget have not gone unnoticed by policymakers, and the ACA will dramatically reduce government payments to plans. These reductions in plan payments might be expected to lead some firms to withdraw from Medicare, and possibly affect firms' decisions to offer PDP plans. Unlike Medicare Advantage plans, PDPs operate under a payment system with a national payment rate determined by the average of PDP plans' bids. However, if Medicare Advantage and PDP plans are demand substitutes, withdrawal of Medicare Advantage plans could induce an increase in the number of PDP plans offered.

After reviewing the relevant literature, covering the necessary background, and providing a conceptual orientation in the next section, we describe our methods, data, and results. The conclusion discusses the implications for market strategy and Medicare payment policy, as well as limitations and extensions of our work.

Background and Conceptual Approach

Entry and Market Structure

The industrial organization (IO) literature characterizes the various roles that market structure plays in market entry decisions. High market concentration is consistent with entry barriers, which can take many forms and arise in multiple ways. In health insurance markets, foreclosure—the reduction in competition due to vertical behavior—can arise from vertical integration (e.g., a hospital-insurer merger), exclusive contracting (e.g., contracts between insurers and providers or advertisers), or most-favored nation requirements (e.g., where an insurer pays unit prices that are no higher than those paid by any other insurer doing business with a hospital) (Gaynor and Vogt 2000).

Vertical behavior has been found to reduce the foreclosing firm's costs and raise those of rivals (Krattenmaker and Salop 1986). Raising rivals' costs has advantages over predatory price reductions, since it does not require "deep pockets" or entail lower (or negative) short-term profits (Salop and Scheffman 1983). Hence, one might expect insurers to enter into exclusive or long-term contracts with lower-cost providers, preventing existing market participants and potential market entrants from doing so. Gal-Or (1996) showed that a provider will accept an exclusive deal with an insurer, even at a lower rate of payment, in return for a larger volume of patients. Encinosa (1996) considered exclusive contracts between physician practices and HMOs. A risk-averse, incumbent HMO may foreclose rivals by entering into an exclusive deal with the only available provider. Gaynor and Vogt (2000) noted, however, that exclusive contracts "appear to be relatively rare between insurers and health care providers [though] long term services contracts are common, and may confer a degree of exclusivity on an insurer who is a large buyer" due to capacity constraints. For years, the American Medical Association has suggested in its annual report on competition in health insurance markets that incumbent insurers' exclusive provider networks reduce competition by making it more costly for potential rivals to enter (AMA 2010). Of the three plan types we consider, only CCPs establish provider networks. Thus, network-based entry barriers are directly possible only for that product type. However, as we will show, there may be spillovers to the PFFS and PDP markets.

Contestability theory suggests that high sunk costs—costs that already have been incurred and cannot be recovered—can also deter market entry if firms anticipate they will serve as an exit barrier (Baumol, Panzer, and Willig 1982). Hilliard, Ghosh, and Santerre (2010) included the costs of marketing and establishing provider networks among insurers' sunk costs. Another barrier to entry is the high switching costs of consumers (Samuelson and Zeckhauser 1988). Age is negatively correlated with health plan switching (Atherly, Florence, and Thorpe 2005), so Medicare beneficiaries are likely "sticky," reluctant to change health insurance plans even if better options exist (Abaluck and Gruber 2009).

Approaches to Modeling Market Structure

The economics literature includes two fundamental approaches to capture effects of competition in models of entry and other market outcomes. Recent work on health care and insurance markets have included both structural (Maruyama 2011; Starc 2010; Lustig 2010) and reduced-form (Dafny and Duggan 2009; Dafny 2010; Bates and Santerre 2008; Schneider et al. 2008; Shen, Wu, and Melnick 2010; Moriya, Vogt, and Gaynor 2010) models. Structural models of entry have been applied in health care (most recently by Maruyama 2011) and, for decades, to problems in nonhealth industries as well (e.g., Berry 1992; Seim 2006). Many of the reduced-form models employ the Herfindahl-Hirschman index (HHI) as an independent variable (Dafny and Duggan 2009; Bates and Santerre 2008; Schneider et al. 2008; Shen, Wu, and Melnick 2010; Moriya, Vogt, and Gaynor 2010), as we do in our application. Though admittedly ad hoc, Gaynor and Town (2011) wrote that "one can think of [such models] as attempting to capture the impacts of relative bargaining power on price, using buyer and seller HHIs as proxies for bargaining power."

Reduced-form models that include HHIs distinguish themselves from structural models in other ways. A reduced-form approach permits the researcher to be agnostic about the underlying competitive game and, thereby, to avoid any game-theoretic assumptions (Gaynor and Town 2011). The trade-off is that reduced-form models do not estimate fundamental parameters associated with a game, as structural models do. A consequence of this distinction is that less precise insight might be gained from a reduced-form model than from a more detailed structural one, but with the advantages of requiring weaker assumptions and leading to easier interpretability and relying on simpler econometric methodology. In addition, despite its shortcomings, the HHI remains an important market measure for policy. The antitrust agencies still use it to inform their analysis of markets for anticompetitive mergers (U.S. DoJ and FTC 2010).

Though we acknowledge that there is an ongoing debate within IO and across empirical economics about the strengths and limitations of structural and reduced-form models (Angrist and Pischke 2010; Nevo and Whinston 2010), we argue that a reduced-form approach is suitable for our application. Outside of health care, structural models have been used to examine a problem similar to one we address: entry decisions by firms that can offer more than one product type. In a modification of the work of Bresnahan and Reiss (1991), Mazzeo (2002) did the path-breaking work, examining entry into motel markets by firms endogenously choosing high, medium, or low quality. Dranove, Gron, and Mazzeo (2003) applied the framework to commercial HMO entry, distinguishing between local and national products. Unfortunately, this approach requires firms to choose only one product type in each market and the approach becomes intractable with more than three types. Although we have only three types (CCP, PDP, and PFFS), firms may enter with one of seven configurations (CCP only, PDP only, PFFS only, CCP-PDP, CCP-PFFS, PFFS-PDP, or CCP-PDP-PFFS). Furthermore, one of our central research interests is to investigate how entry with one product type affects a firm's decision to enter with another product type, and the structural approach studies only the number of entrants of each type. Firm effects within markets cannot be measured.

The Medicare Private Plan Market

Since the emergence of PFFS plans in 2001 and PDPs in 2006, entry and marketing decisions related to those products, as well as enrollment options for beneficiaries, have become more complex. With this additional complexity comes opportunity. Firms may target their entry and marketing efforts geographically, as they did with provider network-based products under Medicare Advantage and its predecessor programs. In addition, they now may differentially allocate their marketing and advertising resources across product types and geography, emphasizing particular plan types where it is more profitable to do so. Moreover, firms are permitted to enroll beneficiaries simultaneously in a PDP and a PFFS product (but in no other combination of products), which motivates the joint marketing of at least those two types of plans.

In one of the few papers that qualitatively describes firm decision making in post-MMA Medicare, Gold (2007) reports on interviews conducted in spring 2006 with representatives of firms offering Medicare products. Firm representatives acknowledged several elements of strategy relevant to our work: 1) firms entered and marketed PFFS plans differentially across counties depending on perceived profitability; and 2) firms adopted a strategy of enrolling beneficiaries into PFFS plans with the goal of moving them to other, more profitable products. The first point, though focused on PFFS plans, is consistent with the notion that firms make different entry and marketing decisions across geographic areas and product types. It is also relevant to our definition of product entry, which requires a signal in the form of nontrivial enrollment that the product is genuinely available to beneficiaries (described in detail later). The second point—that PFFS plans served as feeders to other plan types—is evidence that firms make joint entry and marketing decisions across product types, suggesting that economies of scope may be important and, therefore, that market structure for one product (e.g., CCPs) may be relevant to entry for other products (e.g., PDPs and PFFS plans).

Medicare private plan entry has grown following recent increases in payment (Frakt, Pizer, and Feldman 2009). Payment increases mandated by the MMA and prior legislation have pushed payments to Medicare

Advantage plans well above average per beneficiary costs for traditional fee-for-service Medicare. The average Medicare Advantage payment was 114% of average FFS expenditures in 2009 (Medicare Payment Advisory Commission 2009). In response, the ACA will reduce Medicare Advantage payment rates and use the savings to subsidize insurance for low-income, nonelderly individuals (Kaiser Family Foundation 2010; Biles and Arnold 2010).

The literature on private plan participation in Medicare focuses predominantly on HMO entry in response to government payments and, in many cases, on units of analysis other than the firm. Yet, the conceptual model underpinning the Medicare entry studies assumes that firms enter if expected variable profit exceeds fixed costs. Expected variable profit consists of the government payment rate and the premium collected per enrollee, net of per enrollee costs, multiplied by expected demand. Given a model like this, the literature seeks to measure the relationship between government payments and Medicare entry using an assortment of variables to control for the potentially confounding effects of variations in per enrollee costs and expected demand.

Notably, no prior studies have estimated the effects of market structure on firms' decisions to enter the Medicare market with more than one of the currently permitted product types. Most prior work is based on data from the period before PDPs (and in some cases, PFFS plans) existed (Adamache and Rossiter 1986; Pai and Clement 1999; Porell and Wallack 1990). Because firms can now make entry decisions across all products simultaneously, there is the potential for joint decision making within firms and competition across firms that would be revealed only by a multiproduct entry model such as the one we have developed.

Except for recent work on PFFS entry (Frakt, Pizer, and Feldman 2009), the unit of analysis in Medicare private plan entry studies is the county (Cawley, Chernew, and McLaughlin 2005), contract (Gold 2005), or plan (Abraham et al. 2000). Only one type of plan is analyzed or all plan types are grouped together (i.e., a firm enters Medicare if it enters with any plan type). In the earliest estimate of the effect of government payments on Medicare participation by HMOs, Adamache and Rossiter (1986) found that increasing the payment rate $10 above the national average increased the probability of entry by 2.7 percentage points. Porell and Wallack (1990) found that high payment levels, prior Medicare contract experience, and prior federal qualification were the most important factors distinguishing market participants from nonparticipants. Pai and Clement (1999) reported that the effect of a one-standard-deviation increase in payment rates on the probability of market participation declined from .22 percentage points in 1986 to .03 percentage points in 1994. Finally, Abraham et al. (2000) found that a one-standard-deviation increase in payment rates was associated with an increase of .043 percentage points in the probability of HMO participation in Medicare over the period 1990 to 1995.

A more recent approach to estimating Medicare private plan participation focuses on market factors that make the county more or less profitable for HMOs to enter. Because the county is the unit of observation, this approach does not allow characteristics of specific plans to appear in the entry model. White and Doksum (2001) found that counties in the highest payment quartile were three to 10 times more likely to have any plan entry from 1994 to 1997, compared with counties in the lowest payment quartile. Cawley, Chernew, and McLaughlin (2002) studied how the number of HMOs participating in a county in each year from 1993 to 2000 varied with the payment level, following an approach similar to Bresnahan and Reiss (1991). In a later analysis, the same authors (2005) estimated the level of government payments necessary to induce various levels of HMO participation in counties from 1993 to 2001. They found 381 counties were paid more than necessary to induce participation by one plan, while 1,463 counties were paid less than the one-plan threshold.

Liu and Town (2003) estimated a model at the county-plan level in which the dependent variable was whether at least one Medicare HMO was available in the county in each year from 1993 to 2000. Liu (2003) investigated the determinants of Medicare HMO availability in urban counties and in two categories of rural counties. Simulation analysis showed that using payments as the sole tool to improve rural areas' access to Medicare HMOs could be ineffective or too expensive.

Maruyama (2011) investigated the welfare effects of the consequences of various Medicare HMO payment policies for entry. The structural modeling framework captured product differentiation and firm heterogeneity and allowed for endogenous prices and entry. In addition to being a reduced-form approach, our work differs from that of Maruyama in several other important respects. Maruyama studied the entry of only one product type, HMOs, while our interest is on firm decisions across three products: CCPs, PDPs, and PFFS plans. Additionally Maruyama's unit of observation was the plan-year-county, while ours is a firm-level analysis.

Implications of New Policies for Empirical Analysis

Though considerable attention has been paid to the recent expansion and cost of private plans in Medicare, it is less widely recognized that most of the growth is concentrated in a relatively small number of firms. Firms can expand participation in Medicare in three ways: 1) by increasing the geographic extent of entry; 2) by increasing the number of plans of a given type offered within a geographic region; and 3) by increasing the number of plan types offered. The first of these takes advantage of economies of scale (e.g., of centralized administrative services), whereas the latter two leverage economies of scope. Neither of the existing approaches to the study of market entry—the firm's decision to participate in Medicare as a whole or the presence of one private plan type at the county level—can account for the multifaceted nature of Medicare private plan participation as it exists today.

In contrast to the prior literature that focuses on just one product type, our concern is the firm's decision to enter Medicare with multiple products. This invites the consideration of inter-product competition among firms and strategic marketing of multiple products by a single firm. Broadly, there are two dimensions of inter-product competition among firms. One involves offering drug benefits. All three plan types can compete by offering drug coverage, though among the Medicare Advantage plan types only the CCP is required to offer at least one plan design with a bundled drug benefit. Another dimension is non-drug benefits, within which only CCPs and PFFS plans can compete. Our estimates most strongly shine light on competition between PDPs and the other two plan types.

With respect to strategic marketing across plan types within the same firm, the firm's advertising efforts are likely to be shared inputs. Sharing these efforts across product types would increase public awareness of all the firm's products, resulting in each dollar of advertising reaching more people. Sales forces also can realize economies of scope by representing more than one product to customers. As described previously, there is anecdotal evidence that firms adopt exactly these types of strategies in Medicare (Gold 2007). Moreover, PFFS plans and PDPs can be jointly purchased, possibly adding to the efficiency of the joint marketing of those two plan types. To the extent economies of scope are facilitated by market power, they may be more readily realized by larger firms in concentrated markets (Demsetz 1973). Prior work by Engberg et al. (2004) finds little evidence of scale and scope economies among commercial, Medicaid, and Medicare HMOs. However, our analysis is sensitive to scope economies among different plan types within the Medicare market (i.e., plans that market to the same population), where they more plausibly may exist.

Methods

We estimated two types of models: 1) independent univariate probits for entry by each private plan type, and 2) a multivariate probit entry model that allows for correlated error terms in the individual entry equations. The year-county-firm is the unit of observation for both models. Each observation includes one entry indicator for each of the three plan types (CCPs, PDPs, and PFFS plans). The entry indicators are the dependent variables in the entry equations. The explanatory variables include measures of product-level market structure (instrumented as subsequently explained), Medicare payments to Medicare Advantage plans, plan cost and demand variables, as well as year, firm, and census division fixed effects.

One of the principal independent variables is product-level market structure. Berry and Reiss (2007) have illustrated the role of market structure in firms' entry decisions. A potential entrant anticipates how other firms will react to its entry and from this expectation decides whether it can charge a profitable price. Amel and Liang (1997) postulated that entry is a function of the difference between the firm's expected profit and the level of entry-deterring profits. The former depends on expected post-entry market concentration. They used the three-year mean of the HHI as a proxy for expected market structure.

Likewise, we postulate that a potential entrant into any of the three Medicare product markets anticipates the post-entry market structure. Following Amel and Liang (1997), we proxy expected market structure with a measure of observed structure: contemporaneous product-level HHI. Consistent with the recent work on the effect of commercial health insurance market structure on premiums (Dafny, Duggan, and Ramanarayanan 2009) and the quantity of hospital services used (Bates and Santerre 2008), we account for the endogeneity of market structure with a two-stage residual inclusion (2SRI) instrumental variables approach (Terza, Basu, and Rathouz 2008; Terza, Bradford, and Dismuke 2008), computing standard errors via bootstrapping (Efron 1979).

Our preferred specification includes only the CCP HHI. Barriers to entry arising from market concentration are most likely to occur in the market for CCPs, which raise competitors' costs by tying up providers explicitly through contracting or other vertical behavior or implicitly via capacity constraints, as discussed earlier. To the extent entry is correlated across products within firms, barriers to CCP entry should also affect PDP and PFFS entry. Our estimates reveal the extent to which this is the case.

Market structure for other plan types also could be important for joint entry because the three plan types compete along specific dimensions. However, we did not include market structure for PDPs and PFFS plans in the preferred specification because these plan types do not establish hospital and physician networks, so the connection between concentration and entry barriers is less likely. In particular, entry into the PDP sector has been robust, with about 50 plans available in every market area during our period of study (Frakt and Pizer 2006). We estimated, but do not show, versions of the joint entry models that included the other HHIs in addition to the CCP HHI. The HHIs for non-CCP products were not statistically significant. Our preferred specification reveals the only robust results.

Our instruments for the current CCP HHI are historical HHIs from other segments of the health insurance market, the Federal Employees Health Benefits Program (FEHBP) and commercial group products. This approach is similar to Santerre and Vernon (2007/2008) and Grabowski and Hirth (2003), who instrumented for nonprofit nursing home market share with lagged market share of another service, nonprofit hospitals. In our case, lagged FEHBP and commercial group health insurance HHIs are plausible instruments for current Medicare CCP HHIs because they identify markets more favorable to higher (or lower) concentration based on persistent characteristics of provider organizations or populations not otherwise captured by observable controls. The instruments are also plausibly valid (excludable from the second stage) because FEHBP and commercial insurance products are not substitutes or complements for Medicare products.

Using historical FEHBP and commercial group insurance HHIs at the state level as instruments further decouples them from current, county-level Medicare product entry decisions. The state-level values, in general, are not highly influenced by any particular, single county-level market. In that sense, they are close to approaches taken in other applications where a key product characteristic—usually price—is instrumented with the average price from neighboring markets within the same firm (Hausman, Leonard, and Zona 1994; Nevo 2001; Town and Liu 2003; Frakt and Pizer 2010). Our instrument is constructed from neighboring- and same-market data within each state out of necessity. The data available to us, described in the Data section, includes only state- and firm-level enrollment by market type (such as FEHBP or commercial), so we could not exclude values in a firm's own county-level market. We report tests of instrument strength and excludability in the Results section.

Another variable key to firms' entry decisions in Medicare is the county-level Medicare Advantage payment rate. This is most plausibly relevant for Medicare Advantage plans and perhaps less so for PDPs, which are not paid at Medicare Advantage rates but instead receive a payment from Medicare that does not vary geographically. For this reason, and the fact that our model includes year effects, Medicare payments to Part D plans cannot be included in the model. Yet, as we show, the Medicare Advantage payment rate is an important predictor of PDP entry as well. Our measure of the payment rate is the county Medicare Advantage "benchmark." Though not exactly the same thing as the actual payment rates—which are unavailable to researchers—benchmarks are highly correlated with payments. Since 2006, Medicare Advantage plans have been paid no more than an administratively set benchmark. Plans that bid below the benchmark are paid the benchmark less 25% of the benchmark-bid difference (MedPAC 2009).

Following the standard in the literature, we assume Medicare Advantage payment rates are exogenous (Chernew, DeCicca, and Town 2008). In particular, it is plausible to assume they are not correlated with unobservable cost factors. Though private Medicare plan payments were once directly related to lagged FFS cost, statutory adjustments to the payment rates have severed this link. The 1997 Balanced Budget Act reduced payments overall while increasing them for plans in rural areas with low FFS costs. The Medicare, Medicaid, and SCHIP Benefits Improvement and Protection Act of 2000 extended the payment increases to low-cost urban areas as well. The 2003 MMA introduced additional changes to the payment formula so that by 2005 Medicare Advantage payment rates were about 15% above average fee-for-service costs (Biles et al. 2006; Frakt, Pizer, and Feldman 2009). Beneficiary premiums and benefits are influenced by changes in payment rates, which would not happen if payments were just tracking plans' costs (Pizer, Frakt, and Feldman 2003).

Plan cost variables at the county level include per capita FFS cost, the proportion of elderly age 75 years or older, doctors and hospital beds per capita, urban/rural county indicators, Medigap premiums, and the "risk score," a diagnosis-based measure of average health status (Pope et al. 2004).

Demand variables at the county level include per capita income and the proportions of the population who are elderly, in poverty, have a high school diploma, have four or more years of college, and work in manufacturing, construction, or white-collar jobs. These labor force variables are significant predictors of Medicare plan entry (Cawley, Chernew, and McLaughlin 2005; Pizer, Feldman, and Frakt 2005). All variables are listed with univariate descriptive statistics in Table 3.

All models were estimated with Stata 10 (StataCorp 2007). The trivariate probit model was estimated using methods developed by Cappellari and Jenkins (2003, 2006). The equations are:

$$y_1 = \beta_{10}\text{HHI} + \beta_{11}\text{benchmark} + \beta_{12}\text{cost} + \beta_{13}\text{demand} + \beta_{14}\text{FE} + \gamma_1 + \varepsilon_1 \tag{1}$$

$$y_2 = \beta_{20}\text{HHI} + \beta_{21}\text{benchmark} + \beta_{22}\text{cost} + \beta_{23}\text{demand} + \beta_{24}\text{FE} + \gamma_2 + \varepsilon_2 \tag{2}$$

$$y_3 = \beta_{30}\text{HHI} + \beta_{31}\text{benchmark} + \beta_{32}\text{cost} + \beta_{33}\text{demand} + \beta_{34}\text{FE} + \gamma_3 + \varepsilon_3 \tag{3}$$

where CCP entry occurs if $y_1 > 0$, PDP entry occurs if $y_2 > 0$, and PFFS entry occurs if $y_3 > 0$. The β coefficients are vectors of appropriate size that take different values across the three equations. The fixed effects

(FE) include year, firm, and census division. In particular, the firm fixed effects control for correlations of entry across markets. One instance of each type is omitted so constant terms (γ) are included and vary across equations. The error terms ε have zero mean and unit variance. In the trivariate probit version, the error terms are correlated across equations.

HHI is a two-element vector. The first element is the year-state CCP HHI, computed as the sum of squared market shares for CCPs; it is the dependent variable in a first-stage ordinary least squares (OLS) regression that includes all control variables from the second stage and two additional identifying instruments: state-level HHIs for FEHBP and group markets in 2005. The second element of HHI is the residual from the first-stage regression (Pizer 2009). Standard errors for the second-stage univariate probits are computed by bootstrapping (Efron 1979). Only small deviations between bootstrapped and non-bootstrapped standard errors were observed; none that affected qualitative conclusions. The computing time required for trivariate probit estimation prohibited computation of bootstrapped standard errors for that model. There were only small differences between the coefficients of the independent univariate models and the trivariate model.

Data

We constructed an analytical file for the period 2007–2009 with the year-county-firm as the unit of observation. Each observation includes separate indicators for CCP, PDP, and PFFS entry. The principal building blocks of the data file are geographic service area, firm/plan characteristics, beneficiary counts, and enrollment data sets available from the Centers for Medicare and Medicaid Services (CMS) for each year from 2007 to 2009. To these we merged additional CMS information and secondary data from a variety of sources, described subsequently.

The most reliable sources of plans' service areas are mid-year versions of the Prescription Drug Plan Formulary and Pharmacy Network Files[5] (for PDPs) and the Medicare Options Compare (MOC) database[6] (for CCPs and PFFS plans), which we obtained for the 2007–2009 period. We merged enrollment data at various levels of aggregation. CMS provides files with county-plan-level enrollment for 2008 and 2009 but not 2007. Therefore, we estimated county-plan-level enrollment for 2007 by raking, an iterative post-stratification method used in survey research to adjust weights to meet a set of known partial sums (Lohr 1999). We had CMS data for the partial sums of county-plan-level enrollment over counties within plans (giving national plan totals) and, separately, over plans within a county contract (giving county contract totals). By iteratively matching each set of partial sums, the raking estimator converges to county and plan enrollment estimates for 2007. To test the estimation accuracy, we also computed raking estimates for 2008 for which we had known county-plan enrollment values. A regression of the 2008 estimates on the known values had an R^2 of .95.

We also merged county-level Medicare beneficiary counts obtained from Medicare Advantage market penetration files for all three years.[7] Following Frakt, Pizer, and Feldman (2009), we computed and merged county-level Medicare Advantage benchmarks and FFS costs for all years. The CMS diagnosis-based risk score for 2006,[8] Area Resource File (ARF)[9] variables from various years prior to our study window, and Medigap premiums for 2005 from a large insurer were all merged at the county level. Group and FEHBP HHI instruments were computed from 2005 state- and firm-level covered lives data purchased from the National Association of Insurance Commissioners (NAIC).[10] California and Mississippi did not report FEHBP enrollment data to the NAIC, so observations for those states were excluded from the analysis.

Excluding the small number of records with missing or inconsistent merge keys or other data, the analytic file included almost 700,000 year-county-plan records distributed nearly evenly across the period 2007–2009. As expected, given the relative levels of participation in Medicare, 6% of records were CCPs, 65% were PDPs, and 29% were PFFS plans.

In contrast to some prior studies, we distinguish between *contracting* and *entry* with a meaningful level of enrollment. A contract is an agreement between a firm and CMS that permits the firm to offer a plan type (CCP, PDP, or PFFS) in a geographic area in the contract year. Although some studies use contracting as the signal of entry, contracting could occur without a meaningful level of enrollment. We followed Cawley, Chernew and McLaughlin (2005) and Maruyama (2011) and dropped all plans with very low enrollment. One interpretation of very small enrollment is that it signals that the plan's parent firm is not actively marketing the plan, and that the firm has not entered the market in the same sense as with other plans with nontrivial enrollment. Thus, we dropped all plans with enrollment either below 11 (CMS's cutoff for enrollment reporting) or with market share below .05%. A substantial amount of contracting is not associated with a meaningful level of enrollment. About 43% of county contract pairs account for only 1% of total enrollment in our data. With the remaining data we defined entry for each plan type as occurring when a firm entered a particular county with that plan type in a specific year (using a meaningful level of enrollment as described previously). Note that this only defines where and when a firm and plan type have entered, but does not completely define where and when the firm did not enter. For the latter, we need to specify the set of counties where the firm may feasibly enter.

It is not reasonable to view a county as being in a firm's feasible entry set if the firm faces large barriers to entry in that county, apart from those related to market structure. We think the most important of such non-market structure barriers are state licensing requirements (Horoschak and Silva 2007). For instance, there is no risk that Blue Cross and Blue Shield of Michigan will enter counties in California because it is not licensed to do business in that state. Because we do not have data on these types of entry barriers by state and firm we defined the feasible entry set as follows: a firm is at risk for entering every county in a state with all plan types in all years if and only if it entered at least one county in that state with any plan type in any year. Thus, a firm that entered just one county in state *S* with just one product in just one year is at risk for entering all counties in state *S* with all three products in all three years. Maruyama employs a similar notion of "potential entrants," requiring potential market participants to have some activity in a nearby market.

Of the 199 unique firms represented by our data, 151 (75%) offered only one type of plan (CCP, PDP, or PFFS plan) across the study period and across all counties. Such firms are not our focus because their entry behavior does not reflect multiproduct entry decisions and can be modeled with relatively simple techniques (see White and Doksum 2001 for an HMO entry model and Frakt, Pizer, and Feldman 2009 for a PFFS entry model). Thirty-two firms (16%) offered two types of plans across the study period and across all counties, though not necessarily in the same location at the same time. These firms also are not our focus, although their behavior could be modeled as a bivariate joint decision. Our focus is on 16 firms (8%) that offered all three plan types, again not necessarily contemporaneously or in the same location. While this is a small proportion of all Medicare-participating firms, it reflects 89% of PDP enrollment, 84% of PFFS enrollment, and half of CCP enrollment. Nevertheless, our model, and the inferences based upon it, are not generalizable to all firms participating in Medicare, only to the firms that have demonstrated a willingness to offer all three plan types according to the criteria noted previously. The 16 firms in our sample represent 83,577 year-county-firm multiproduct entry decisions.

Results

Table 1 describes the entry patterns of the 16 firms in our data. Most firms in the data are national players; six are only active in a small minority of counties. The first column reports the percentage of counties at risk for entry. By the definition in the previous section, only firms that never enter a state with any product are considered to have no risk for entry in the counties of that state. However, once a firm has entered any county in a state, in any year and with just one product, we

Table 1. Risk set and entry patterns by firm and product types, 2007–2009

Firm	Percent of counties at risk for entry	Percent of counties in risk set entered with:							
		None	CCP only	PDP only	PFFS only	CCP and PDP only	CCP and PFFS only	PDP and PFFS only	All
All firms in sample	61.3	21.1	.1	50.8	.4	2.3	.0	23.6	1.7
Aetna	100.0	36.4	.0	58.4	.2	2.6	.0	2.2	.1
BCBS of Florida	2.4	1.0	.0	62.7	.0	16.4	.0	18.9	1.0
BCBS of Michigan	2.9	.0	.0	.0	.8	.0	.0	64.7	34.5
BCBS of South Carolina	7.2	14.1	.0	38.9	4.7	.0	.0	38.9	3.4
Bravo Health	55.6	64.7	.2	34.0	.0	.7	.0	.3	.0
Cigna	100.0	34.8	.0	64.4	.0	.1	.0	.7	.0
Coventry Health Care	100.0	11.7	.0	41.0	1.3	.7	.0	43.5	1.9
Geisinger Health System	4.3	77.0	10.7	1.6	1.6	2.5	5.7	.3	.5
Health Net	100.0	56.3	.0	39.0	.6	.7	.0	3.3	.1
Highmark	4.3	.8	.0	35.8	.0	57.7	.0	4.9	.8
Humana	100.0	.4	.0	22.7	.0	1.4	.0	71.5	3.9
UnitedHealth Group	100.0	1.4	.0	60.7	.1	7.3	.0	23.2	7.3
Universal American	100.0	2.5	.0	54.8	.1	.4	.0	41.4	.7
University of Pittsburgh Medical Center	4.3	55.7	6.8	24.0	.0	12.3	.0	.8	.3
Wellcare Health Plans	100.0	9.5	.0	69.2	.1	2.4	.0	17.9	.8
Wellpoint	100.0	10.9	.0	62.5	.8	3.2	.0	21.9	.7

Notes: Table is based on year-county-firm study data; $N = 83{,}577$. The number of counties at risk for entry can be computed by multiplying the percent of counties at risk by the number of year and county pairs (8,520).
BCBS= Blue Cross Blue Shield; CCP= coordinated care plan; PDP=prescription drug plan; PFFS=private fee-for-service plan.

Table 2. Premium and cost-sharing, by plan type, 2009

	Plan type					
	CCP		PDP		PFFS	
	Percent with no premium	Mean for plans with premiums ($)	Percent with no premium	Mean for plans with premiums ($)	Percent with no premium	Mean for plans with premiums ($)
Nondrug premium (per month)	67.5	68			39.4	56
Primary care doctor visit costs						
Deductible	99.9	68			99.2	121
Copayment	32.9	11			2.9	21
Drug costs (for plans offering drug benefits)						
Drug premium (per month)	65.8	25	.0	35	17.0	30
Deductible	89.8	283	58.4	280	66.4	233
Mean preferred Rx copay	.0	33	.0	37	.0	35

Notes: N = 71,836 county-plan records.
All figures were weighted by county-plan enrollment. CCP= coordinated care plan; PDP= prescription drug plan; PFFS= private fee-for-service plan.

consider all counties in that state at risk for entry by that firm in all years (including prior ones) and for all products.

Each cell in the other columns of Table 1 reports the proportion of counties in the entry risk set entered by a given firm with the indicated plan type(s), aggregated across all three years, 2007–2009. The categories in a row are mutually exclusive and collectively exhaustive (cells pertaining to product entry rates in a row sum to 100%). The first row is aggregated across all firms and illustrates three predominant entry modalities: firms do not offer any products in about 21% of counties; they offer only a PDP in about 50% of counties; and they offer a PDP and a PFFS plan in about 23% of counties. All other combinations of plan type entry are far less common.

However, there is considerable variation in entry patterns by firm and product. Nine firms (Aetna, Cigna, Coventry, Health Net, Humana, UnitedHealth Group, Universal American, Wellcare, and Wellpoint) are "national firms," in the sense that their entry risk set comprises 100% of U.S. counties. Of the national firms, Humana has the largest entry footprint, having entered 99.6% of the counties in its risk set.

We also observe variation in entry strategies. Certain firms have clearly adopted a strategy that differs from the average results in the first row. For example, Aetna, Bravo Health, Cigna, Health Net, Highmark, and the University of Pittsburgh Medical Center infrequently offer PFFS products; Blue Cross Blue Shield of Michigan offers some type of product in every county in its risk set; and Geisinger is much more concentrated in the CCP market than other firms. Overall, the results of Table 1 suggest firms make strategic multiproduct entry decisions that vary a great deal over different plan types and firms.

Table 2 illustrates the degree of beneficiary cost-sharing for all county and plan pairs associated with the 16 firms in Table 1. The 2009 nondrug plan premium and cost-sharing (deductible and copayment) for primary care doctor visits and the drug premium, deductible, and mean preferred drug copayment are summarized. Each cell includes the percentage with no premium or cost-sharing and the mean for those paying some amount. All figures are weighted by county-plan-level enrollment.

The results are consistent with summaries of benefits published elsewhere using different samples, methods, and years (Gold 2005, 2007, 2009; Hargrave et al. 2009; Kaiser Family Foundation 2009). CCP enrollees are more likely than those in PFFS plans not to

Table 3. Variable definitions and univariate statistics

Variable	Description	Mean	[Min, Max]	Source
Entry				
CCP entry	CCP entry indicator, 2007–2009	.042 (.20)	[.00, 1.00]	CMS
PDP entry	PDP entry indicator, 2007–2009	.78 (.41)	[.00, 1.00]	CMS
PFFS entry	PFFS entry indicator, 2007–2009	.26 (.44)	[.00, 1.00]	CMS
Market structure				
CCP HHI	State-level CCP HHI, 2007–2009	.36 (.20)	[.093, 1.00]	CMS
Grp. mkt. HHI	State-level group market HHI, 2005	.30 (.18)	[.081, .91]	NAIC
FEHBP HHI	State-level FEHBP HHI , 2005	.70 (.25)	[.19, 1.00]	NAIC
Payment				
Benchmark	Benchmark payment rate, 2007–2009	769 (74)	[692, 1307]	CMS
Cost				
FFS cost	Average FFS cost, 2007–2009	657 (80)	[436, 1285]	CMS
Prop. elderly 75+	Proportion of elderly 75+ years old, 2000	.47 (.045)	[.21, .62]	CMS
Docs. per capita	General practitioners per capita, 2006	.031 (.021)	[.00, .26]	ARF
Hosp. beds per capita	Hospital beds per capita, 2005	.0035 (.0053)	[.00, .11]	ARF
Rural county	Rural county, 2003	.29 (.45)	[.00, 1.00]	ARF
Urban county	Urban county, 2003	.36 (.48)	[.00, 1.00]	ARF
Rx Medigap prem.	Monthly drug Medigap premium, 2005	238 (34)	[188, 466]	-[a]
Non-Rx Medigap prem.	Monthly nondrug Medigap premium, 2005	137 (26)	[103, 263]	-[a]
Risk score	Aged/disabled risk score, 2006	.97 (.069)	[.70, 1.35]	CMS
Demand				
Prop. eld. in poverty	Proportion elderly in poverty, 1999	.12 (.056)	[.00, .48]	ARF
Per capita income	Per capita income in thousands, 2005	26.95 (6.90)	[.00, 93.4]	ARF
Prop. HS diploma	Proportion of population age 25+ with a high school diploma, 2000	.77 (.087)	[.35, .97]	ARF
Prop. 4+ yrs. col.	Proportion of population age 25 with 4+ years college, 2000	.16 (.078)	[.049, .64]	ARF
Prop. manufacturing	Proportion of workers in manufacturing, 2000	.16 (.089)	[.003, .48]	ARF
Prop. white collar	Proportion of white-collar workers, 2000	.51 (.090)	[.31, .84]	ARF
Prop. construction	Proportion of workers in construction, 2000	.078 (.02)	[.017, .23]	ARF

Notes: Table based on year-county-firm study data. $N = 83{,}577$. Year, firm, and state fixed effects are not shown. All variables are at the county level except where indicated. Standard deviations are in parentheses. CCP= coordinated care plan; PDP= prescription drug plan; PFFS= private fee-for-service plan; HHI= Herfindahl-Hirschman index; FEHBP= Federal Employees Health Benefits Program; CMS= Centers for Medicare and Medicaid Services; NAIC= National Association of Insurance Commissioners; ARF=Area Resource File.

[a] Provided by a large insurer.

have a nondrug premium, but when there is a cost, it is $12 higher on average. Enrollees in both CCPs and PFFS plans almost always have a zero deductible for primary care doctor visits. Those in CCPs are more likely to have a zero copayment for a doctor visit than PFFS plans; when they have a copayment it is $10 lower. Enrollees in drug-offering CCPs pay the lowest drug premium, are most likely to have no drug deductible, and have the lowest mean preferred drug copayment. PDP enrollees pay the highest drug premium, are least likely to have a zero drug deductible, and face the highest mean preferred drug copayment.

Table 3 provides definitions and descriptive statistics for key variables in the models of product entry (year, firm, and census

Table 4. Estimation results: first-stage CCP HHI ordinary least squares

Variable	Coefficient
Instruments	
Grp. mkt. HHI	−.082 (.0043) ***
FEHBP HHI	.47 (.0037) ***
Payment	
Benchmark[a]	−.00078 (.0010)
Cost	
FFS cost[a]	−.0021(.00096) **
Prop. elderly 75+	−.055 (.014) ***
Docs. per capita[b]	.18 (.025) ***
Hosp. beds per capita	2.25 (.10) ***
Rural county	.026 (.0013) ***
Urban county	.010 (.0014) ***
Rx Medigap prem.	−.0010 (.000027) ***
Non-Rx Medigap prem.	.00057 (.000034) ***
Risk score	−.31 (.0094) ***
Demand	
Prop. eld. in poverty	.32 (.014) ***
Per capita income	.00088 (.000097) ***
Prop. HS diploma	−.12 (.013) ***
Prop. 4+ yrs. col.	−.23 (.014) ***
Prop. manufacturing	.15 (.0085) ***
Prop. white collar	.42 (.016) ***
Prop. construction	−.17 (.025) ***
	R^2 = .54
	Instruments' F-statistic = 11,610

Notes: Constant and year, firm, and census division fixed effects are not shown. Table is based on year-county-firm study data. N = 83,577. Standard errors are in parentheses. CCP= coordinated care plan; HHI= Herfindahl-Hirschman index; FEHBP= Federal Employees Health Benefits Program.
[a] Benchmark and FFS cost variables have been divided by 100.
[b] Docs. per capita variable has been multiplied by 100.
* Significant at the .1 level; ** significant at the .01 level; *** significant at the .001 level.

division fixed effects are not shown). Variables are organized by category of principal relevance: entry (dependent variables), market structure (endogenous CCP HHI and instruments for it), payment, cost, and demand. Some variables differ by year, as indicated in the table. All except market structure vary at the county level. County-level CCP HHIs cannot be computed because many counties have no entrants; HHIs are undefined in such cases. Moreover, instruments were available only at the state level.

Table 4 presents the coefficients for the first-stage OLS regression of CCP HHIs. FEHBP HHIs are positively correlated with the CCP HHI. Group market HHI is negatively correlated with CCP HHI, perhaps because Medicare CCPs are not group products and are subject to different market dynamics. The coefficient on benchmark payment rates is not statistically significant. All other coefficients are statistically significant. Note that the benchmark and FFS cost variables have been divided by 100 and doctors per capita has been multiplied by 100.

Our instruments have an F-statistic well above the standard threshold of 10 (Staiger and Stock 1997) and far exceeding the Stock-Yogo 10% critical value for maximal size (Stock and Yogo 2005) indicating that they are sufficiently strong. We also performed Sargan (1958) and Basmann (1960) tests of overidentification using linearized probability model (two-stage least squares) versions of the models shown in Table 5 (discussed later). Though these are typically discussed as tests of excludability, they are, in fact, joint tests of excludability *and* homogeneity of treatment effects (Angrist 2010). Consequently, instruments that are excludable may be rejected

Table 5. Estimation results: second-stage independent univariate probit

Variable	CCP entry		PDP entry		PFFS entry	
	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect
Market structure						
CCP HHI	−1.40 (.14) ***	−.33	−.53 (.092) ***	−.11	−.58 (.085) ***	−.14
CCP HHI 1st stage residual	−.38 (.15)**	−.092	.10 (.10)	.021	.14 (.096)	.033
Payment						
Benchmark[a]	.27 (.019) ***	.065	−.18 (.014) ***	−.04	.41 (.014) ***	.10
Cost						
FFS cost[a]	−.017 (.021)	−.0041	.094 (.013) ***	.020	−.79 (.013) ***	−.19
Prop. elderly 75+	−1.15 (.31) ***	−.28	−.83 (.19) ***	−.17	−1.52 (.18) ***	−.36
Docs. per capita[b]	−2.09 (.68) **	−.50	3.14 (.34) ***	.65	3.49 (.32) ***	.83
Hosp. beds per capita	−6.14 (2.64) *	−1.47	−13.87 (1.33) ***	−2.87	.65 (1.34)	.15
Rural county	−.59 (.049) ***	−.14	−.33 (.017) ***	−.069	−.28 (.017) ***	−.067
Urban county	.38 (.027) ***	.090	.29 (.020) ***	.061	.21 (.017) ***	.050
Rx Medigap prem.	−.0027 (.00042) ***	−.00065	−.0032 (.00042) ***	−.00065	−.0025 (.00045) ***	−.00060
Non-Rx Medigap prem.	.0026 (.00077) ***	.00063	.0040 (.00049) ***	.00083	.0027 (.00052) ***	.00064
Risk score	2.56 (.22) ***	.61	2.89 (.14) ***	.60	.026 (.13)	.0063
Demand						
Prop. eld. in poverty	−2.65 (.38) ***	−.64	−2.50 (.20) ***	−.52	−.63 (.21) **	−.15
Per capita income	.0061 (.0017) **	.0015	.0083 (.0016) ***	.0017	.0024 (.0013) *	.00057
Prop. HS diploma	−1.31 (.29) ***	−.31	−2.97 (.18) ***	−.62	−.90 (.17) ***	−.21
Prop. 4+ yrs. col.	−1.50 (.28) ***	−.36	.59 (.20) **	.12	−.65 (.18) ***	−.16
Prop. manufacturing	1.98 (.19) ***	.47	3.96 (.11) ***	.82	3.68 (.11) ***	.87
Prop. white collar	5.69 (.36) ***	1.36	4.69 (.25) ***	.97	3.30 (.21) ***	.78
Prop. construction	−.38 (.59)	−.09	1.61 (.34) ***	.33	1.16 (.34) ***	.28
Pseudo R^2	.35		.44		.40	

Notes: Constant and year, firm, and census division fixed effects are not shown.Table is based on year-county-firm study data. $N = 83{,}577$. Marginal effects were computed at the mean. Standard errors are in parentheses. CCP= coordinated care plan; PDP= prescription drug plan; PFFS= private fee-for-service plan; HHI= Herfindahl-Hirschman index.

[a] Benchmark and FFS cost variables have been divided by 100.

[b] Docs. per capita variable has been multiplied by 100.

* Significant at the .1 level; ** significant at the .01 level; *** significant at the .001 level.

due to local average treatment effects. The p-values for the null hypothesis of joint excludability and effects homogeneity for our two additional instruments (FEHBP and group product HHIs) are .07, .32, and .34, for the CCP, PDP, and PFFS equations, respectively.

Table 5 provides coefficient estimates and marginal effects for the second-stage univariate probit models of product entry, estimated by 2SRI with bootstrapped standard errors. Estimates for the CCP HHI coefficients are statistically significant and negative, consistent with the presence of barriers to entry. Incumbent CCPs may use exclusive contracts with providers or other vertical restrictions to reduce the likelihood of entry by potential rivals. If entry in a county with a CCP is unattractive for a firm, that firm may be less likely to enter with other products because it is less able to leverage economies of scope. This may be why CCP market concentration is a barrier for PDP and PFFS plan types as well.

As expected, higher benchmarks are associated with a higher likelihood of CCP and PFFS entry. That they are associated with lower PDP entry may be due to cross-product competition. That is, counties more conducive to CCP and PFFS entry (those with higher benchmarks) are likely to present firms with lower residual demand for PDPs, discouraging entry with a PDP product. Also, to the extent that CCP and PFFS plans experience favorable selection, PDPs may wish to avoid the remaining high-risk beneficiaries.

Signs of most other coefficients are as expected. Taking the cost-related explanatory variables first, higher FFS cost reduces entry of CCP and PFFS plans (though in the CCP case, the coefficient is not statistically significant), but increases entry of PDPs. The CCP and PFFS results are consistent with lower profit in higher-cost counties. The PDP result again may be due to cross-product competition. FFS Medicare costs are likely less predictive of PDP costs than CCP and PFFS costs. In areas less attractive to CCPs and PFFS plans, PDPs have less competition (both from other firms and within the same firm). And again, if CCPs and PFFS plans would otherwise enroll the good risks, areas where they are absent present opportunities for PDPs to experience more favorable selection.

Entry in counties with a higher proportion of elderly over age 75 is less attractive for all plan types, possibly due to higher risk that is not completely built into the risk-adjusted payment formula. More doctors per capita are associated with greater probabilities of entry for PDPs and PFFS plans. More hospital beds per capita are associated with lower probabilities of entry (where significant). Firms are relatively more likely to offer PFFS plans in rural counties and CCPs in urban counties. This is consistent with the costs associated with establishing provider networks. Higher Medigap drug premiums deter entry of all product types, perhaps reflecting cost factors not accounted for in FFS costs, which are based only on non-drug utilization. The non-drug Medigap premium is positively associated with entry of all product types and may reflect aspects of demand not captured by other demand effects (described later). Firms are more willing to offer products in counties with higher risk scores. This may indicate systematic overpayment in the risk adjustment formulas, especially if firms can manage and/or code to cause beneficiaries to appear higher risk (Angeles and Park 2009; Miller and Luft 1994).

Turning next to the demand-related explanatory variables, demand for private Medicare plans is lower in counties with a higher proportion of elderly in poverty and greater in counties with higher per capita income. Less-educated populations are generally less likely to purchase private Medicare plans, especially PDPs. The labor force variables may reflect firms' decisions in the employer market and heavily unionized industries and occupations, which traditionally offer retiree health benefits to employees. Prior studies have found these variables to be positively associated with CCP entry (Cawley, Chernew, and McLaughlin 2005), and we find the same positive effects for the other plan types.

We also estimated a trivariate probit model that permits cross-product residual correlations.

Table 6. Simulated changes in the probability of market entry in response to payment cut, by plan type, 2009

Plan type	Baseline (%)	Simulated (after payment cut) (%)	Absolute change	Relative change (%)
CCP	13.2	8.0	−5.2	−40.0
PDP	89.2	91.1	1.9	3.1
PFFS	29.6	20.5	−9.1	−30.7

Notes: All results are beneficiary weighted. CCP= coordinated care plan; PDP= prescription drug plan; PFFS= private fee-for-service plan.

These correlations represent the effects of unmeasured cost and demand variables that are common to plan types in the same firm. The highest residual correlation is between PDP and PFFS plans, .41, and the lowest is between CCP and PFFS plans, .075. A correlation of .32 was found between CCPs and PDPs. All residual correlations are statistically significant at the $P = .0001$ level. Economies of scope from shared advertising and marketing costs could explain the overall positive pattern of the correlations. For example, if a firm invests in a local advertising campaign creating positive associations with the firm's name, the expected profit from each product should increase, raising the probability of entry with all products simultaneously. That these effects are strongest for the PDP and PFFS combination is consistent with the fact that these products can be bundled together and sold to the same beneficiary, reducing sales costs. While economies of scope are a plausible explanation for these results, they are not the only one. We cannot rule out the possibility that these effects result exclusively from correlated demand shocks instead of cost savings.

The coefficient estimates and standard errors (not shown) of key variables in the trivariate entry model, including HHI, benchmark, and FFS costs, are almost identical to those presented in Table 5 and described previously. This implies that despite highly significant cross-equation correlations between residuals, the independent equations can be used for policy analysis.

To further illustrate the effect of the relationship between plan entry and payment, we used the coefficients reported in Tables 4 and 5 to simulate the ACA provisions that will reduce Medicare Advantage payments. The law will gradually lower Medicare Advantage benchmarks to levels that are multiples of per beneficiary FFS costs, fully phasing in for all plans in 2016 (Biles and Arnold 2010). The new benchmarks will range from 95% of per beneficiary costs in counties in the top quartile of FFS costs to 115% of per beneficiary costs for counties in the bottom quartile of FFS costs. New plan payments will be capped at old payment rate levels so that the new benchmarks can only lower, not raise, payments. Our simulation of payment rate changes incorporates the aforementioned features of the ACA provisions. We did not simulate other features, like the proportion of benchmark bid differences that plans can retain for below-benchmark bids (currently 75%, this will be reduced to as low as 50%, depending on plan quality measures) and quality-based bonus payments (Biles and Arnold 2010).

The simulation proceeded as follows. First, using the models of Table 4 we simulated the effect of the reduced benchmarks on the CCP HHI. Next, we used the models of Table 5 to predict the probability of entry of each plan type in each county in 2009, based on new payments and the HHIs they are predicted to induce. The simulation results are shown in Table 6 as absolute and relative changes in the predicted probability of plan entry from baseline (beneficiary weighted). Consistent with the estimated model coefficients of Table 5, the probabilities of CCP and PFFS availability decline considerably, while that for PDPs modestly increases.

Conclusion

We developed firm-level models of product entry into the Medicare private plan market. The models differ from previous ones in that they treated as distinct the three principal plan types (CCPs, PDPs, and PFFS plans), and were estimated at the level at which joint decisions are made—the firm. We also estimated a model that allowed for correlated errors among the entry equations.

We found that higher CCP market concentration reduces the probabilities of entry for all plan types, consistent with the hypotheses of entry barriers and economies of scope. An entry barrier for CCPs could be the establishment of exclusive provider networks. The effects of CCP market concentration on the entry of other plan types can be explained by economies of scope. Because entry is positively correlated across products within firms, entry barriers that apply to CCP plans also deter entry of other plan types. The relevance of this finding for antitrust policy is that CCP market concentration affects market structure across the entire spectrum of Medicare product markets, not just within the CCP market.

Another important policy implication of our findings is that proposed reductions in Medicare Advantage payment rates are likely to increase market entry by firms with PDP products and decrease entry of CCPs and PFFS plans. These effects are shown both in our estimated coefficients and in a simulation of the scheduled payment cuts. The PDP response is consistent with strategic competition. CCPs and PFFS plans are less likely to enter counties where they are paid less, all other things equal. These counties offer opportunities for PDPs due to lower competition in the dimension of drug benefits. To the extent that CCPs and PFFS plans experience favorable selection, counties where they are absent offer PDPs a more favorable risk pool, all other things equal. This finding, while readily interpretable as evidence of inter-firm competition, can also result from intra-firm strategy. Even within a single firm, counties relatively less favorable for CCP and PFFS entry could be more favorable for PDP entry for the reasons already given.

We also found significant cross-product residual correlations—highest between PDP and PFFS plans, and lowest between CCP and PFFS plans. These correlations reflect unmeasured cost and demand variables common to plan types within the same firm. While such results are consistent with economies of scope (e.g., from shared advertising and marketing costs), we cannot identify such effects apart from other unmeasured characteristics. However, the fact that the largest positive correlation is found between PDP and PFFS plans—the only plan types that permit joint enrollment—is further suggestive of economies of scope. The small positive correlation found between CCP and PFFS plans could be due to the unmeasured cost of developing a provider network that applies only to CCP plans.

Last, we provided descriptive evidence that highlights dramatic differences in strategy across firms, with some focusing on local marketing of a single product and others marketing all three products across practically the entire country.

Our work has several limitations. First, our model does not necessarily generalize to all firms participating in Medicare. To study multiproduct entry across three plan types, we restricted the sample to firms that had offered all three products, though not necessarily contemporaneously and in the same location. However, firms offering all three plan types enroll most Medicare beneficiaries in private plans. Simpler models can be estimated for firms that only offer one or two products. Additionally, we could not consider competition from Medigap plans because we are not aware of Medigap data comparable to that available for CCPs, PDPs, and PFFS plans. Second, we do not directly model economies of scope among Medicare private plan types, although we suggest they are present and cautiously interpret the positive cross-equation correlations as indicative of their effects. We lack direct measures of plan costs that would be necessary to estimate a multiproduct cost function.[11] Third, we do not attempt to assess the welfare implications of a cut in Medicare Advantage payment rates such as the one we simulate. Finally, as with any empirical model, ours is based on historical data. As such, this imposes limitations on the extent to which we can accurately predict the future with our model-based simulation. We cannot address potential changes that might occur and which are outside the range of our data.

Despite these limitations, this paper is the first to examine the effects of market structure in the three main Medicare product markets and to illustrate firms' multiproduct entry behavior in the current Medicare policy regime. As additional experience is gained and data are collected over a wider range of payment rates, models like those estimated in this paper can be used to support a wider range of policy simulations.

Notes

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1 It is also relevant that one cannot model PDP entry at the county level aggregated over firms, since PDPs have entered all counties.
2 Part D is Medicare's prescription drug program.
3 Part C is Medicare's program for private comprehensive health plans that bundle the benefits of Part A (inpatient hospital insurance), Part B (insurance for physician and outpatient services), and, optionally, Part D (drug insurance). In addition to CCPs and PFFS plans, Medicare Part C also includes other plan types with very small enrollment: regional PPOs (3% of Medicare Advantage enrollment), medical savings accounts (1,866 enrollees), and other plan types and demonstrations (collectively accounting for 3% of Medicare Advantage enrollment) (Henry J. Kaiser Family Foundation 2009).
4 Most PFFS plans have only an implied network consisting of providers willing to accept it. However, even if a provider accepts a PFFS patient, the provider is under no obligation to do so for subsequent visits. In contrast, CCPs establish contractual relationships with providers. Inclusion in a CCP's network is contingent upon acceptance of specified payment arrangements (Blum, Brown, and Frieder 2007).
5 Prescription Drug Plan Formulary and Pharmacy Network Files are available for order at: http://www.cms.hhs.gov/NonIdentifiableDataFiles/09_PrescriptionDrugPlanFormulary,PharmacyNetwork,andPricingInformationFiles.asp.
6 MOC data are downloadable from http://www.medicare.gov/download/downloaddb.asp.
7 Market penetration files are downloadable from http://www.cms.hhs.gov/MCRAdvPartDEnrolData/. Because market penetration files are not available for 2006 or 2007 we used December 2005 data for 2007.
8 The risk score is included in CMS's FFS data: http://www.cms.hhs.gov/MedicareAdvtgSpecRateStats/05_FFS_Data.asp.
9 The ARF can be obtained from the U.S. Department of Health and Human Services: http://www.arf.hrsa.gov.
10 The National Association of Insurance Commissioners (NAIC) does not endorse any analysis or conclusions based upon the use of its data.
11 Engberg et al. (2004) estimated a multiproduct cost function for a related set of products: commercial HMOs, Medicare HMOs, and Medicaid HMOs.

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