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Informative Censoring by Health Plan Disenrollment among Commercially-insured Adults

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Abstract

Purpose: Health plan disenrollment occurs frequently in commercial insurance claims databases. If individuals who disenroll are different from those who remain enrolled, informative censoring may bias descriptive statistics as well as estimates of causal effect. We explored whether patterns of disenrollment varied by patient or health plan characteristics.

Methods: In a large cohort of commercially-insured adults (2007–2013), we examined two primary outcomes: a) within-year disenrollment between January 1 to December 30, which was considered to occur due to patient disenrollment from the health plan; and b) end-of-year disenrollment on December 31, which was considered to occur due to either patient disenrollment from the health plan or withdrawal of the entire health plan from the commercial insurance database. In yearly cohorts, we identified factors independently associated with disenrollment by using log-binomial regression models to estimate risk ratios (RR) and 95% confidence intervals (CI).

Results: Among 2,053,100 unique patient-years, the annual proportion of within-year disenrollment remained steady across years (range, 13% to 14%) whereas the annual proportion of end-of-year disenrollment varied widely (range, 8% to 26%). Independent predictors of within-year disenrollment were related to health status, including age, comorbidities, frailty,

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CONFLICT OF INTEREST

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hospitalization, emergency room visits, use of durable medical equipment, use of preventive care, and use of prescription medications. In contrast, independent predictors of end-of-year disenrollment were related to health plan characteristics including insurance plan type and geographic characteristics.

Conclusions: Differential risk of disenrollment suggests that analytic approaches to address selection bias should be considered in studies using commercial insurance databases.

Keywords

administrative claims; epidemiologic methods; informative censoring; pharmacoepidemiology; selection bias

INTRODUCTION

Healthcare insurance claims databases are used in a variety of research settings to study population-level use of therapeutics and associated outcomes.¹ These data sources permit the study of entire populations and provide longitudinal follow-up on individuals. However, in commercially-insured populations in the US, health plan disenrollment is thought to be common and leads to censored observation of outcomes.

Commonly used approaches for censored data, such as Kaplan-Meier estimators of survival functions, assume that censoring is independent from the outcome under study. However, it is possible that individuals who disenroll are different in a multitude of ways compared to those who remain enrolled. If the predictors of disenrollment are not independent of the study's outcome, then the censoring (i.e., loss of follow-up data) after disenrollment is 'informative' and the use of many standard methods may bias descriptive statistics as well as estimates of causal effect.^{2,3} In this context, appropriate approaches to address selection bias should be used, such as inverse probability-of-censoring weighted estimators or regression adjustment.⁴

Even though disenrollment in commercial insurance claims databases is thought to be ubiquitous, the assumption that disenrollment is unrelated to prognostic patient factors has not been formally examined in the epidemiologic literature. In the present study, we explored whether patterns of health plan disenrollment varied by patient- or health plan-level characteristics in a large cohort of commercially-insured adults in the United States. We sought to examine temporal trends and predictors of disenrollment. We also explored whether results were modified by the date of disenrollment.

METHODS

Data source and study population

We used the Truven Health MarketScan Commercial Database® for the years 2007–2013 (Copyright 2013 Truven Health Analytics Inc. All Rights Reserved). These databases contain de-identified individual-level information on health insurance enrollment, inpatient and outpatient diagnoses and procedures, and outpatient pharmacy-dispensed medications for employer-based commercial insurance beneficiaries, their spouses, and dependents.⁵ We

used a retrospective cohort design and constructed six yearly cohorts (2008–2013). To create a study population large enough to achieve expected precision estimates, each yearly cohort was constructed as a 2% random sample of individuals enrolled during that cohort year. We required adults to be enrolled in health plans on January 1 of each year as well as a 6-month baseline period of continuous enrollment and prescription drug coverage from July 1 to December 31 of the preceding year. Follow-up spanned January 1 to December 31 of each year. Individuals were eligible for multiple yearly cohorts. To eliminate the impact of Medicare enrollment among older adults, we restricted the analyses to adults less than 64 years of age on January 1st; therefore, individuals were not eligible to experience disenrollment from their commercial insurance during follow-up due to Medicare enrollment on his/her 65th birthday. We did not attempt to exclude individuals with disabilities who may qualify for Medicare coverage (e.g., end-stage renal disease) given the expected low proportion within this commercially-insured population.

Outcome definitions

Information on date of disenrollment was identified from the enrollment detail table. Disenrollment from the health plan was defined as the first date of a >7-day gap in coverage. We examined two primary outcomes: a) within-year disenrollment between January 1 to December 30, which was considered to occur due to patient disenrollment from the health plan; and b) end-of-year disenrollment on December 31, which was considered to occur due to either patient disenrollment from the health plan or withdrawal of the entire health plan from the commercial insurance database.⁶ Patient disenrollment from the health plan could be due to death, which was not well-captured in these data. Individuals who changed health insurance plan type but did not accrue a gap in coverage >7 days were not considered to have disenrolled. For the analysis of within-year disenrollment, follow-up began on January 1 and continued until the earliest date of disenrollment or end of the follow-up period on December 30. For the analysis of end-of-year disenrollment on December 31, the cohort was restricted to those still enrolled on December 30. As a secondary analysis, we defined disenrollment from the health plan using a larger gap in coverage (>30-day).

Covariates

Baseline covariates were ascertained during the 6-month baseline period. Demographic characteristics included age on January 1st, sex, geographic region, and residence in a metropolitan area (defined by the U.S. Census Bureau as a geographic region with a high population density at its core and close economic ties throughout the area).⁷ Relationship of the individual to the primary beneficiary was categorized as employee, spouse, or child/other (hereforth, child). Health care utilization variables, which served as proxy metrics to identify individuals in worse health, included overnight hospitalizations and emergency department visits. Therapeutic classes of prescription drugs, based on categories defined by IBM RED BOOK®, were included in the model as indicator variables based on frequency of use in the study population. In addition to selecting commonly used drugs, we also selected drugs thought to potentially be related to disenrollment or death (e.g., prescriptions associated with abuse or serious comorbidities). Comorbidities, frailty markers, and healthcare utilization variables were defined using *International Classification of Diseases, Ninth Revision, Clinical Modification* (ICD-9-CM) diagnosis and procedure codes and Health Care

Financing Administration Common Procedure Coding System (HCPCS) codes (Table S1 in supplementary material).

Statistical analysis

We calculated descriptive statistics of patient characteristics. To understand temporal trends in disenrollment, we calculated monthly and annual proportions by timing of disenrollment. Using log-binomial regression models to estimate risk ratios (RR) and 95% confidence intervals (CI), we identified factors independently associated with health plan disenrollment. Separate models were created for two outcomes: a) within-year disenrollment; and b) end-of-year disenrollment (restricted to individuals still eligible on December 30). We used the robust variance estimator to allow for clustering at the individual level. Analyses were further stratified by year. We assessed the discriminative ability of the model to predict disenrollment using logistic regression to estimate the C-statistic. To remove the influence of dependent children, a secondary analysis was performed excluding adults <27 years of age. All analyses were performed using SAS 9.3 (SAS Institute, Cary, NC, USA) and R statistical software 3.3 (Vienna, Austria).⁸ This study was approved by the Institutional Review Board at Washington University School of Medicine.

RESULTS

We identified 2,000,842 eligible individuals contributing person-time during 2,053,100 unique patient-years. The study population had a median age of 44 years, 47% were male, and 84% resided in a metropolitan statistical area. During the 6-month baseline period, 8% had an emergency room visit, 3% had an overnight hospitalization, and 64% filled prescription medications for 1 therapeutic classes (Table 1). Table S2 provides additional characteristics of the study sample. Disenrollment occurred frequently, ranging from 22% in 2008 to 40% in 2012. The annual proportion of within-year disenrollment remained steady across years (range, 13% to 14%), whereas the annual proportion of end-of-year disenrollment varied widely across years (range, 8% to 26%) (Figure 1). The monthly proportion of within-year disenrollment remained steady from January to November (0.84% to 1.75%); in contrast, within-year disenrollment appeared very low in December (range, 0.01% to 0.03%) because disenrollment usually occurred on the last day of the month (i.e., end-of-year disenrollment) (Figure S1). In the secondary analysis varying the definition of disenrollment from a coverage gap of >7 to >30 days, the proportion of individuals who disenrolled within-year decreased by 0.7% and end-of-year by 0.1%.

The adjusted RRs and 95% CIs for the relationships between patient characteristics and disenrollment, by timing of disenrollment, are presented in Figure 2 and Table S3. Age was independently associated with both within-year and end-of-year disenrollment, with stronger age effects associated with within-year disenrollment. Compared to individuals 30–39 years of age, within-year disenrollment was less likely among individuals who were older and more likely among individuals who were younger (with the exception of 18–21 year olds). The associations for within-year and end-of-year disenrollment followed similar patterns in secondary analyses restricted to adults 27–63 years of age (Table S4). Year-stratified results indicated large changes over time in risk of within-year disenrollment for adults 18–26 years

of age (Figure S2). In 2012 and 2013, the risk of within-year disenrollment was dramatically lower among 18–24 year olds, whereas the risk of within-year disenrollment was dramatically higher among 25–26 year olds. In addition, beneficiary status was independently associated with both within-year and end-of-year disenrollment. Compared to male employees, dependent children were more likely to disenroll within-year and end-of-year while female spouses were less likely to disenroll within-year.

Both geographic region and health insurance plan type were independently associated with disenrollment at any time, though associations were stronger for end-of-year disenrollment than for within-year disenrollment (Figure 2, Table S3). Individuals living in the Midwest generally were least likely to disenroll within-year but most likely to disenroll at end-of-year. Individuals on a health plan type of preferred provider organization or exclusive provider organization were most likely to disenroll, with much stronger associations between health insurance plan type and end-of-year disenrollment than within-year disenrollment.

Several comorbidities and frailty markers were associated with elevated risk of within-year disenrollment, including but not limited to cancer, coagulopathy, stroke, HIV, myocardial infarction, and psychiatric diagnoses (Figure 2, Table S3). Some of these conditions were also associated with end-of-year disenrollment, however, associations were generally attenuated to the null.

Healthcare utilization associated with preventive care was commonly associated with lower risk of within-year disenrollment (Figure 3, Table S3). For example, within-year disenrollment was less likely among individuals who received influenza vaccination (RR 0.92, 95% CI: 0.91, 0.93), lipid testing (RR 0.95, 95% CI: 0.94, 0.96), an eye exam (RR 0.92, 95% CI: 0.90, 0.93), a screening pelvic exam (RR 0.95, 95% CI: 0.92, 0.97), a screening prostate-specific antigen test (RR 0.93, 95% CI: 0.91, 0.95), or a screening mammogram (RR 0.89, 95% CI: 0.88, 0.91). One exception to this pattern was pneumococcal vaccination was associated with higher risk of within-year disenrollment. For end-of-year disenrollment, there was more variability in the associations related to preventive care.

Healthcare utilization associated with disease or frailty, including use of durable medical equipment, was predictive of within-year disenrollment, but rarely associated with end-of-year disenrollment (Figure 3, Table S3). For example, individuals with recent emergency department visits had higher risk of within-year disenrollment (RR 1.14, 95% CI: 1.12, 1.15) but not end-of-year disenrollment. Similarly, individuals with recent overnight hospitalizations had higher risk of within-year disenrollment (RR 1.09, 95% CI: 1.07, 1.11) but not end-of-year disenrollment. The risk of within-year disenrollment was elevated for individuals who used oxygen equipment (RR 1.26, 95% CI: 1.18, 1.34), a home hospital bed (RR 1.18, 95% CI: 1.02, 1.38), mobility aids (RR 1.11, 95% CI: 1.04, 1.18), or ambulance / life support (RR 1.04, 95% CI: 1.01, 1.08).

The adjusted RRs and 95% CIs for the relationships between prescription drug utilization and disenrollment, by timing of disenrollment, are presented in Figure 4 and Table S3. The

risk of within-year disenrollment decreased with increasing number of therapeutic class prescriptions; the corresponding associations for end-of-year disenrollment followed a similar pattern but were farther from the null. Several therapeutic classes were associated with elevated risk of within-year disenrollment (from strongest to weakest associations): loop diuretics, benzodiazepines, opiate agonists, anti-depressants, amphetamine-type stimulants, and anti-diabetic (insulins). Anti-hyperlipidemic agents and other cardiovascular agents were associated with reduced risk of within-year disenrollment. The associations for end-of-year disenrollment were generally attenuated to the null. The model fit was poor for within-year disenrollment (C-statistic = 0.62) and end-of-year disenrollment (C-statistic = 0.64).

DISCUSSION

In a cohort of commercially-insured adults drawn from a large, administrative claims database, disenrollment from the health plan was independently associated with patient and health plan characteristics, and these associations were modified by the timing of disenrollment. The strongest independent predictors of within-year disenrollment were related to health status, including age, comorbidities, frailty, hospitalization, emergency room visits, use of durable medical equipment, use of preventive care, and use of prescription medications. In contrast, the strongest predictors of end-of-year disenrollment were related to health plan characteristics (type of health insurance plan) and geographic characteristics (region and rural-urban status). Disenrollment was a common event, ranging from 22% in 2008 to 40% in 2012. Over the six-year study period, the annual proportion of within-year disenrollment remained constant whereas the annual proportion of end-of-year disenrollment fluctuated extensively.

Age was the strongest independent predictor of within-year disenrollment, with substantially higher disenrollment among younger adults, particularly those age 22–24 in 2008–2011, and those age 24–26. These findings may be partially due to early career job transitions. In addition, younger adults may disenroll from a family health plan on which they are listed as a dependent either because they start a job through which they get their own health insurance or because they age out of eligibility as a dependent; the beneficiary status variable in our predictive models indicated that adult children covered by the family health plan were more likely to disenroll than employees or spouses of the same age. In 2011, mid-way through our study period, the Affordable Care Act (ACA) implemented a requirement that all plans and issuers offer dependent child coverage to make coverage available until a child reaches 26 years of age.⁹ The impact of the ACA legislation is evident in our results, where 18 to 24 year olds were more likely to experience within-year disenrollment from 2008–2010, and less likely to experience within-year disenrollment from 2011–2013. Those age 25–26 had high rates of within-year disenrollment across the entire study period, but with the highest rates from 2011–2013, when the age limit for dependents was 26 years.

After accounting for age, within-year disenrollment was more likely among individuals with markers of poor health and less likely among users of preventive health services. Individuals most likely to disenroll within-year included those with comorbidities; emergency department visits and overnight hospitalizations; use of durable medical equipment; and

prescriptions for several types of therapeutic classes. Conversely, individuals were less likely to disenroll within-year if they had recently received cancer screenings; an eye exam; lipid testing; antihyperlipidemic agents; or influenza vaccination. Receipt of pneumococcal vaccination was the only type of preventive care associated with increased risk of within-year disenrollment. However, in our study population of adults <65 years, pneumococcal vaccination is likely a marker of poor health rather than preventive care because pneumococcal vaccination is only recommended for adults 19–64 years with immunocompromising conditions, functional or anatomic asplenia, cerebrospinal fluid leak, or cochlear implants.^{10,11} Apart from young adults who frequently disenroll from health plans, our results indicate that individuals who remained enrolled appear to be healthier and more health-seeking than individuals who disenroll. It is plausible that individuals in poor health are less capable of maintaining employment and more likely to experience job termination, retirement, or death.

In the epidemiologic literature, factors associated with health plan disenrollment have not been previously explored in commercially-insured adults, but have been identified in Medicare managed care and Medicaid enrollees.^{12–21} Similar to our findings, studies in older adult populations with Medicare insurance generally report that increased frailty and health care utilization are associated with health plan disenrollment from managed care plans. One study of community-dwelling, Medicare managed care enrollees concluded that poor perceived physical and mental health, and a decline in health in the past year independently predicted disenrollment.¹⁸ Another study found that Medicare HMO beneficiaries who disenrolled had 180% the rate of inpatient services in the year after disenrollment as beneficiaries who remained enrolled.²¹ Among beneficiaries in an acute/long-term managed care plan, disenrollment was associated with hospital admission but not age, functional or cognitive impairment, Medicaid eligibility, or diagnoses; these results may be related to the low level of disenrollment in the study population which was characterized by the most frail and vulnerable Medicare beneficiaries.¹⁷ Altogether, these previous findings in the epidemiologic literature support our results of differential disenrollment by patient characteristics - otherwise known to health economists and insurance market analysts as the consequences of “favorable selection” or “adverse selection.”^{22,23}

Our findings of differential risk of disenrollment by patient- and health plan-level characteristics have potentially broad implications for longitudinal studies using healthcare databases. The current standard in epidemiological research of commercially-insured populations is to treat disenrollment as a censoring event that occurs independently of the outcome (noninformative censoring). However, our results raise the possibility that treating health plan disenrollment as an independent censoring event may bias estimates of comparative effectiveness and other parameters of interest. In the context that disenrollment does not occur independently of the outcome (informative censoring), use of standard methods may yield estimates subject to selection bias. The potential for selection bias will depend on the amount of disenrollment and the strength of its association with the exposure and the outcome under study. In scenarios with high potential for selection bias, standard methods may be biased, as demonstrated by previous theoretical and applied research.^{2,3,24–27} In this context, appropriate methods to address selection bias should be promoted, including but not limited to stratification-based techniques including standard regression

adjustment, joint modeling, and inverse probability of censoring weighted estimation.
3,4,28–30

Our findings are subject to several limitations. First, we made simplifying assumptions about the underlying causes and timing of disenrollment in order to isolate factors that drive disenrollment. We assumed that patient disenrollment from the health plan occurred throughout the year, whereas withdrawal of the entire health plan from the commercial insurance database only occurred on December 31st. In practice, it may be possible that withdrawal of entire health plans from the commercial insurance database could occur prior to December 31st. However, our assumption probably holds because a higher proportion of individuals disenrolled on December 31st (14.4%) than from January 1st to December 30th combined (13.8%). Even if this assumption holds, we are unable to distinguish the type of end-of-year disenrollment; individuals may disenroll on December 31st due to withdrawal of the entire health plan from the commercial insurance database or patient disenrollment from the health plan. Nevertheless, our findings of dissimilar predictors of within-year versus end-of-year disenrollment support the hypothesis that disenrollment on December 31st was generally due to withdrawal of the entire health plan from the commercial insurance database.

Second, commercial claims insurance databases do not capture the reason for disenrollment. Our attempt to characterize factors associated with patient disenrollment from the health plan portrays a heterogeneous group of disenrollees; within-year disenrollment appears to occur frequently among: 1) young healthy individuals leaving the family plan or experiencing a change in employment; and 2) individuals in poor or declining health with high healthcare utilization who likely experience job termination, early retirement, or death. The MarketScan database captures inpatient deaths but not outpatient deaths, which precludes the ability to fully incorporate mortality information into the analysis. Therefore, analyses of MarketScan data implicitly treat death as a censoring event and thus attempt to estimate the effect of treatment in a setting in which the risk of death is not present (i.e., an estimand that is neither identifiable nor of clinical interest in settings where death is common). In settings where complete death information is available in the data source, death is more appropriately handled as a competing risk.

Finally, we made decisions about the study population and study design that impacted the interpretation of our results. We restricted our study population to adults <65 years, therefore it remains unknown how temporal trends and predictors of disenrollment differ between commercially-insured children, adults 18–64 years, and adults ≥65 years; our results can only be generalized to adults 18–64 years of age. We also created six separate yearly cohorts of randomly sampled individuals and ascertained disenrollment during a one-year follow-up period. This study design enabled characterization of annual trends for within-year and end-of-year disenrollment, but did not permit information on the frequency of disenrollment during longer follow-up periods (as are common in multi-year pharmacoepidemiology studies) or whether the associations between various prognostic factors and disenrollment change over time. Further, our analysis examined the association between health plan disenrollment and factors measured during the 6-month baseline period. The analysis did not incorporate time-varying factors, and thus the health status of enrollees could have improved

or declined during the follow-up period. Some analytic solutions to address selection bias could incorporate time-updated covariates; subsequent analyses could estimate associations between patient- and plan-level characteristics and disenrollment specific to the study sample, research question, and analytic plan of interest.

In this study, we quantify and characterize the common occurrence of health plan disenrollment among adults in a commercial insurance database. By identifying factors associated with health plan disenrollment, we attempted to describe the subpopulation with shorter duration of available follow-up for longitudinal studies using commercial insurance databases. We found frequent disenrollment among two distinct subpopulations: young adults leaving the family plan or experiencing early career job transitions; and individuals in poor or declining health with high healthcare utilization. Otherwise, individuals who remained enrolled in their health plan appear to be systematically healthier and more health-seeking than individuals who disenroll from the health plan. Health plan characteristics, rather than individual level characteristics, were associated with end-of-year disenrollment largely due to withdrawal of the entire health plan from the commercial insurance database. Future research in this area is needed to gain a better understanding of settings under which the treatment of disenrollment as an independent censoring event may induce selection bias and to determine the utility of study design and analytic methods to minimize such bias.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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KEY POINTS:

- Despite the ubiquity of disenrollment in commercial insurance claims databases, the assumption that disenrollment is unrelated to prognostic patient- and health plan-level factors has not been examined in the epidemiologic literature.
- The strongest independent predictors of disenrollment during the year were age, emergency department visits, hospitalizations, comorbidities (e.g., cancer), frailty markers (e.g., difficulty walking), use of durable medical equipment (e.g., oxygen), use of preventive care (e.g., cancer screenings), and use of prescription medications (e.g., loop diuretics).
- In contrast, the strongest predictors of end-of-year disenrollment were health plan type and geographic characteristics.
- Differential disenrollment by patient- and health plan-level characteristics suggests that analytic approaches to address selection bias should be considered in studies using commercial insurance databases.

Figure 1.
Annual trends in disenrollment within a large commercial insurance claims database by timing of disenrollment, United States, 2008–2013.

Figure 2. Adjusted risk ratio and 95% confidence interval estimates for the associations between patient characteristics and health plan disenrollment, United States, 2008–2013.

Abbreviations: CI, confidence interval; COPD, chronic obstructive pulmonary disease; CDHP, Consumer Directed Health Plan; EPO, Exclusive Provider Organization; HIV, human immunodeficiency virus; HMO, Health Maintenance Organization; IQR, interquartile range; POS, Point of Service; PPO, Preferred Provider Organization.

^a Model adjusted for all variables listed as well as healthcare utilization and prescription drug utilization variables presented in Figures 3 and 4, respectively.

Figure 3. Adjusted risk ratio and 95% confidence interval estimates for the associations between healthcare utilization and health plan disenrollment, United States, 2008–2013.

Abbreviations: CI, confidence interval; DME, durable medical equipment; PSA, prostate-specific antigen.

* Model adjusted for all variables listed as well as patient characteristic and prescription drug utilization variables presented in Figures 2 and 4, respectively.

Figure 4. Adjusted risk ratio and 95% confidence interval estimates for the associations between prescription drug utilization and health plan disenrollment, United States, 2008–2013.

Abbreviations: CI, confidence interval.

* Model adjusted for all variables listed as well as patient characteristic and healthcare utilization variables presented in Figures 2 and 3, respectively.

Table 1. Baseline characteristics of 2,053,100 commercially-insured adult patient-years, United States, 2008–2013.^a

Characteristic	Patient-years (n = 2,053,100)	
	N	%
Age, median (interquartile range), years	44 (32, 54)	
Male sex	973,930	47
Relation to employee ^b		
Employee	1,282,510	63
Spouse	553,137	27
Child/other	217,453	11
Region		
Northeast	296,664	15
Midwest	517,397	25
South	800,429	39
West	408,028	20
Missing	30,582	2
Metropolitan statistical area		
No	305,791	15
Yes	1,717,847	84
Missing	29,462	1
Comorbidities		
Hypertension	226,037	11
Type II diabetes	91,443	5
Chronic obstructive pulmonary disease	68,154	3
Cancer	16,723	1
Psychiatric diagnoses	108,333	5
Emergency department visit(s)	172,519	8
Overnight hospitalization(s)	56,254	3
Number of prescription medication therapeutic classes		
0	746,186	36
1–2	559,984	27
3–5	463,325	23
6	283,605	14

^aOf 2,000,842 eligible individuals in the study population, 50,307 (2.5%) individuals were included in two annual cohorts; 950 (0.1%) individuals were included in three annual cohorts; and 17 (0.0%) individuals were included in four annual cohorts. In total, eligible individuals contributed 2,053,100 unique patient-years.

^bStudy population includes employer-based commercial insurance beneficiaries, their spouses, and dependents.