Ordeal Mechanisms, Treatment Continuation, and Health Outcomes: Evidence from Continuous Glucose Monitoring in Medicaid

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Abstract

We quantify the impact of increasing the frequency of insurance reauthorization for previously approved medical treatment on treatment continuation and hospitalizations. We specifically focus on the use of continuous glucose monitoring (CGM) among West Virginia Medicaid beneficiaries who have a Type 1 Diabetes diagnosis. Although CGM is cost-effective, the use of such technology is subject to prior authorization and recurring reauthorization requirements that have become increasingly stringent. Using administrative Medicaid claims data, we find that shortening the renewal interval is associated with substantial reductions in sustained CGM use, with the timing of discontinuation matching the timing of renewals. We also find increases in the probability of hospitalization following a shortening of the re-authorization window. These findings highlight how cost containment policies can generate unintended consequences by restricting access to cost-effective medical care.

Keywords: Prior Authorization, Ordeal Mechanisms, Continuous Glucose Monitoring, Medicaid Policy, Administrative burden.

JEL Classification: I13, I1, H51.

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1 Introduction

Ordeal mechanisms are an often-used tool in government programs, as standard economic theory indicates they can be effective in improving targeting efficiency (Besley and Coate, 1992; Zeckhauser and Nichols, 1982). This theory has recently been called into question, as empirical evidence shows that ordeal mechanisms can decrease targeting efficiency, leading to new models that allow ordeal mechanisms to affect targeting efficiency differently by group type and by the timing of when the ordeal is imposed (Deshpande and Li, 2019; Wu and Meyer, forthcoming). In the context of healthcare, prior authorization is an ordeal mechanism. Prior authorization means that before patients can be authorized to use a certain treatment or receive a specific type of care, a medical provider must provide documentation to the insurance company that convinces the insurance company of medical necessity for that treatment/care. Prior authorizations expire, even for treatments for chronic health conditions in which there is no cure; therefore, even treatment that has passed prior authorization will require additional paperwork in the form of a reauthorization. Failure to successfully complete the reauthorization process results in a denial during treatment, which means that unless the patient is willing and able to pay out of pocket for continued use, it is likely the treatment will be discontinued.

Prior authorization and reauthorization are designed by insurers to contain costs, but they are controversial. Proponents of these policies argue that they improve healthcare by ensuring medical care and treatments are in line with medical standards of care and by shifting patients toward less costly medical care (Americas Health Insurance Providers, 2020). Opponents argue that this bureaucracy causes unnecessary delays and denials (American Medical Association, 2025), empowers insurance company bureaucrats to overrule physicians on medical decisions (Yaver, 2026), negatively impacts patient health resulting in severe adverse events like hospitalization (American Medical Association, 2025; Sausser, 2026), negatively impacts patient's finances and mental health (Sausser, 2026), increases physician burnout (American Medical Association, 2025), and imposes direct costs on doctors requiring them to hire staff entirely dedicated to handling prior authorizations (American Medical Association, 2025). More than half of state legislatures are considering legislation this year that would limit the use of prior authorization (Kirzinger et al., 2026) with people across the political spectrum reporting prior authorization as their single-biggest non-cost concern in accessing healthcare (Sausser, 2026). Recent estimates indicate that each year more than 100 million Americans are impacted by a health insurance denial (Yaver, 2026), and that prior authorizations are becoming more burdensome over time (Kyle et al., 2026; Yaver, 2026). Understanding how renewal requirements affect sustained access to treatment and hospitalizations is thus critical to the debate over prior

authorization policies.

We quantify the impact of changes in the reauthorization window for continuous glucose monitoring (CGM) systems on West Virginia Medicaid beneficiaries with Type 1 Diabetes (T1D). T1D is a chronic and incurable disease that requires monitoring of glucose levels and administering insulin endogenously to keep glucose levels in a healthy range. Severe medical complications, including death, can occur if one's glucose levels are either too high or too low. CGM is a medical innovation that provides real time glucose data and has been shown to improve glycemic control, reduce hypoglycemia, and improve quality of life (American Diabetes Association, 2019; Ni et al., 2023; Prahalad et al., 2020). While clinical guidelines have become increasingly favorable towards patient use of CGM, West Virginia Medicaid twice shortened its CGM reauthorization interval: from one year to six months in the fourth quarter of 2021 and again from six to three months in the second quarter of 2022.[1] These sequential policy changes generate quasi experimental variation in the ordeal mechanism that is independent of underlying clinical need. We can therefore answer the following questions: (1) does this increase in the frequency of reauthorization reduce sustained CGM use among beneficiaries who have already cleared the prior authorization hurdle? and (2) does this increase in the ordeal mechanism increase the probability of an ER visit and/or hospitalization? Using administrative claims data from 2017 through 2022, we construct a panel of adult Medicaid beneficiaries with T1D who have a CGM claim and estimate a dynamic event time model comparing continuation trajectories, ER visits, and hospitalizations across reauthorization regimes within the first year following the first CGM receipt.

The results indicate that shortening the reauthorization interval reduces sustained CGM use. Under the six-month policy, continuation declines more sharply beginning in the second and third quarters after receipt. The reductions are substantially larger under the three-month regime. The timing of these declines closely matches the renewal thresholds, suggesting that increased reauthorization frequency contributes to reduced use over time. These findings are robust to alternative T1D classification algorithms, a model that specifically accounts for the competing risk of attrition from the sample, a comparison of newly initiating versus already-established CGM users, and a placebo test assigning a fictitious reform prior to the policy change.

We also test for heterogeneous treatment effects in CGM continuation by severity of diabetes-related complications. This analysis is motivated by the model developed by Wu

[1] WV is not unique in this respect, as many state Medicaid programs require reauthorization at least annually (Center for Health Care Strategies , 2023). Medicaid has fewer tools at its disposal to contain costs and thus tends to rely on prior authorization more than other types of health insurance.

and Meyer (forthcoming), which harmonizes their empirical evidence that ordeals in the entry phase of public programs increase targeting efficiency while those in the recertification process reduce targeting efficiency. We document significant increases in CGM discontinuation across the distribution of diabetes-related complications, meaning the reauthorization policy did not appear to effectively target the removal of those with the smallest potential for clinical gain from CGM.

Given that previous literature has documented adverse outcomes stemming from discontinuing medical treatment (for example, see Chandra et al. (2024) which finds impacts on mortality), we next test whether there is any impact on the probability of in-patient hospitalization or the probability of an ER visit. We find a significant 10 percentage point increase in the probability of hospitalization three quarters after CGM initiation for those in the six-month reform. If we limit hospitalizations to claims with a T1D reference, there is still a significant increase in hospitalizations (6 percentage points). These results are robust to different T1D definitions. We find no evidence of a significant increase in the probability of an ER visit.

We make five specific contributions to the literature: (1) we isolate the effect of reauthorization frequency, a distinct margin from the presence of prior authorization studied in prior work; (2) we extend the literature on insurance authorization, which has focused on pharmaceuticals, to a medical device with a markedly shorter product life-cycle; (3) we extend the literature on renewals in social programs by focusing on reauthorization for benefits within a program as opposed to recertification to remain part of a program; (4) we bridge the literature on Medicaid administrative burden and the determinants of CGM uptake among publicly insured populations; and (5) we demonstrate the health effects of reducing access to CGM in a real world setting.

This paper contributes to the literature on administrative burdens in healthcare by studying a different margin of administrative burdens; namely, changes in reauthorization intervals. Administrative burdens impact whether people enroll in health insurance (Ericson et al., 2023). They also affect provider behavior, with evidence that complex billing leads to consolidation (League, 2023) while claims denial processes reduce physician participation in Medicaid (Dunn et al., 2024).[2] Other barriers such as preferred drug lists, drug formularies, and prescription drug monitoring programs impose costs on physicians, change physician prescribing behavior, and can even spillover onto prescribing for other patients not directly affected by these policies (Currie et al., 2026; Cutler and Ly, 2011; Epstein and Ketcham, 2014; Ketcham and Epstein, 2008; Wang and Pauly, 2005). While previous work has quantified impacts from adding or removing prior authorization (Alpert and Sun,

[2]See (Currie et al., 2026) for a recent summary of research on doctor decision making.

2026; Burn and Ristovska, 2025; Currie and Musen, 2025; Dillender, 2018; Kyle and Keating, 2024; Mark et al., 2020), our paper adds to the literature by showing that the frequency of reauthorization also matters.

Within the larger literature on ordeal mechanisms in healthcare, we contribute to the literature on whether prior authorization policies achieve their intended policy goals: to reduce costs and better align care with the established standards of care. Interestingly, Shepard and Wagner (2025) point out that despite ordeal mechanisms in health insurance screening out low value consumers, as intended, these individuals tend to also be low cost, raising the question of whether social value outweighs cost in self-screening.[3] For Medicare Part D, using prior authorization leads to net savings (Brot et al., 2023). However, in other contexts, it is argued that the costs of prior authorization exceed the benefits due to treatment discontinuation (Popovian and Winegarden, 2021). When low-cost substitutes exist, prior authorization targeting brand name drugs lead to increased utilization of generic versions (Burn and Ristovska, 2025). Prior authorization rules have also been shown to reduce prescribing of potentially harmful medication (Currie and Musen, 2025; Dillender, 2018). While prior authorization may lead to cost savings, it may come at the expense of some people delaying or not being able to access high-value healthcare (Burn and Ristovska, 2025; Kyle and Keating, 2024; Mark et al., 2020). For example, using variation in age criteria for prior authorization for ADHD medication results in finding reductions in ADHD medication use but not finding improvements in targeting of the medication (Alpert and Sun, 2026). Our paper contributes to the literature by showing that more frequent reauthorization leads to increased treatment discontinuation and increased hospitalizations.

While the related literature has focused on insurance authorizations for pharmaceuticals (Alpert and Sun, 2026; Brot et al., 2023; Burn and Ristovska, 2025; Currie and Musen, 2025; Dillender, 2018; Kyle and Keating, 2024; Mark et al., 2020), our paper studies insurance authorization in the context of medical devices. The regulatory process for medical devices is quite different, and the lifespan of a medical device is much shorter on average than that of a pharmaceutical due to frequent small updates to devices (The Medicare Payment Advisory Commission, 2017). Since prior authorization came about to limit access to new and expensive products (Berg, 2023), and medical devices tend to be newer than pharmaceuticals (The Medicare Payment Advisory Commission, 2017), this suggests that medical devices may be especially impacted by prior authorization rules. With frequent changes to medical devices, manufacturers often make patients switch to the new device by discontinuing old versions, meaning that there is less scope for a medical device equivalent

[3] Incentives may differ in public insurance markets with no or low cost sharing.

to switching to a generic version. For example, one of the two CGM systems covered during the period of this study was the Dexcom G6. Dexcom received FDA approval for the G7 model at the end of 2022, and the product became available in 2023.[4] Dexcom announced the G6 model would stop being manufactured on July 1, 2026.[5] Our paper thus contributes to the literature by studying an ordeal mechanism for medical devices.

The related literature on renewal of social benefits has focused on the margin of recertification to remain in a program, whereas our paper focuses on reauthorization for a benefit within a program. Changes in policy that impact administrative burdens in recertification matter in varied contexts including SNAP (Ganong and Liebman, 2018), Medicaid (Myerson et al., 2025) and financial aid in higher education (Castleman and Page, 2016). Online case management, reminders, and live assistance increased recertification for SNAP and Medicaid (Gray, 2019; Myerson et al., 2025). Automating ex parte Medicaid renewals increases renewals and decreases denials for procedural reasons (Herd et al., 2025). While automation of caseworker assistance improves targeting efficiency at program entry, it reduces targeting efficiency during recertification and leads to more exits (Wu and Meyer, forthcoming). Recertification intervals for social benefits have been specifically noted as an area where more research should be done (Currie, 2006). Kabbani and Wilde (2003) use variation in renewal intervals to document a tradeoff in terms of program participation and error rates in the context of SNAP. Homonoff and Somerville (2021) use variation in timing of the SNAP recertification interview within a month, which affects time available to complete follow up paperwork, to document that the recertification process for SNAP does not improve targeting efficiency. A recent policy report lays out challenges and potential solutions from moving towards social program redeterminations being based on pay period earnings data as opposed to quarterly data (Ganong et al., 2026). Our paper contributes to the literature by documenting that renewal intervals matter not just for program participation, but for continued access to benefits within a social program.

Lastly, much of the prior literature on CGM efficacy focuses on the results of these devices in controlled settings (Foster et al., 2019; Heinemann et al., 2018; Karter et al., 2021; Lind et al., 2017). These RCT results are important and have been instrumental in the development of clinical guidelines supporting expanded CGM use for patients with Type 1 and, increasingly, Type 2 Diabetes (American Diabetes Association, 2019; American Diabetes Association Professional Practice Committee, 2021). However, this is among the

[4]See https://investors.dexcom.com/news/news-details/2022/Dexcom-G7-Receives-FDA-Clearance-The-Most-Accurate-Continuous-Glucose-Monitoring-System-Cleared-in-the-U.S/default.aspx for the announcement

[5]See https://www.dexcom.com/faqs/how-much-longer-will-dexcom-g6-be-available for the announcement.

first studies to demonstrate the real-world improvements in terms of reduced hospitalizations among patients using CGM that uses a quasi-experimental design. This is further supportive evidence of the efficacy of these products and the importance of insurance design in providing access to effective treatments. These patients, due to the chronic nature of T1D and the high costs of hospitalizations associated with complications of ineffective blood sugar management, appear to be the high value consumers that previous literature notes that ordeal mechanisms should not be targeting (Myerson et al., 2025; Wu and Meyer, forthcoming).

The rest of the paper is organized as follows. Section 2 provides background information on T1D, CGM, and Medicaid policy. Section 3 discusses the data and empirical analysis. Section 4 provides the results on how reduced reauthorization windows impact continued CGM usage, ER visits, and hospitalizations. Section 5 provides a discussion and Section 6 concludes.

2 Background on T1D, CGM, and Medicaid Policy

T1D is an auto-immune disease in which the body destroys the cells that produce insulin. People must check their glucose values with a blood glucose meter or through CGM to help determine how much insulin to administer. Both glucose levels that are too high (hyperglycemia) and too low (hypoglycemia) can lead to serious complications or even death if left untreated.[6]

CGM systems are widely recognized as a high-value technology for diabetes management, as they provide frequent updates on glucose levels that can then be used to better calibrate insulin doses. Unlike a glucose meter, CGM does not require a blood sample and instead uses a small medical device that continually provides updated information on glucose values. Medical research has documented substantial evidence of improvements in glycemic control and reductions in hypoglycemia for individuals with T1D which then leads to reductions in short-term costs (Bronstone and Graham, 2016). CGM systems have been demonstrated to be cost-effective (Grunberger et al., 2021; Roze et al., 2020; Wan et al., 2018). The growing evidence on the clinical benefits of CGM has led to changes in the Standards of Diabetes Care, which is provided by the American Diabetes Association. In 2020, it stated that there are benefits for people with T1D "who are not meeting glycemic targets, have hypoglycemia unawareness, and/or have episodes of hypoglycemia" to using CGM (American Diabetes Association, 2019), while in 2022 the recommendation was that

[6]Glucose levels that are too low can lead to diabetic coma, and those that are too high can lead to diabetic ketoacidosis.

CGM should be offered to adults with T1D as long as they are capable of using it (or have caregivers who can) (American Diabetes Association Professional Practice Committee, 2021).

Despite these clinical benefits, CGM adoption remains uneven, particularly among publicly insured populations (Atac et al., 2025; Bailey et al., 2023; DeSalvo et al., 2023; Kruger and Anderson, 2021; Tremblay et al., 2024). A growing body of evidence suggests that insurance design plays an important role in determining access to CGM. Reductions in financial barriers and expansions of Medicaid coverage have consistently increased CGM adoption, improved adherence, and generated better glycemic outcomes among publicly insured patients (Miyazaki et al., 2024; Ni et al., 2023). However, insurance coverage alone does not guarantee access. Prior authorization and reauthorization requirements impose administrative burdens on both beneficiaries and providers that may impede continued use of CGM despite established clinical need (Arbogast et al., 2024; Herd and Moynihan, 2018; Herd and Moynihan, 2025).

Administrative burdens may affect access through both provider and patient channels. In West Virginia Medicaid, beneficiaries treated by board-certified endocrinologists were substantially more likely to adopt CGM following coverage expansion, suggesting that provider expertise and administrative capacity influence access to diabetes technology (Gaulke et al., 2023). More broadly, authorization and billing requirements consume substantial clinical resources, discourage physician participation in Medicaid, and contribute to inefficiencies in healthcare delivery (Casalino et al., 2009; Chandra and Staiger, 2020; Cutler and Ly, 2011; Dunn et al., 2024; Garber and Skinner, 2008; Tanenbaum et al., 2018). Although authorization policies may improve efficiency by reducing unnecessary utilization and healthcare spending (Brot et al., 2023), they may also create barriers to high-value care when repeated documentation requirements interrupt clinically appropriate treatment (Burn and Ristovska, 2025; Kyle and Keating, 2024; Mark et al., 2020; Popovian and Winegarden, 2021). Thus, prior authorization and reauthorization requirements represent a tradeoff between cost containment and access to care. Rather than evaluating the introduction or removal of prior authorization, we isolate the effect of increasing the frequency of reauthorization for beneficiaries who have already demonstrated medical necessity and initiated treatment.

3 Methods

3.1 Data

This study uses administrative claims data from West Virginia Medicaid (WVM) to examine access to CGM systems among adult beneficiaries with T1D. The dataset includes all adjudicated medical and pharmacy claims and allows longitudinal tracking of individuals over time. The study period spans January 2017 (2017-Q1) through December 2022 (2022-Q4). The sample is restricted to beneficiaries aged 20 to 64 years. This age range reflects differences in CGM coverage rules for pediatric beneficiaries and ensures a consistent policy environment across the adult population.[7]

Identification of type 1 diabetes (T1D) is based on claims derived algorithms. In the baseline specification, beneficiaries are classified as having T1D if they have at least one diagnosis of T1D, defined using ICD 10 codes E10.xx. We further restrict the sample to individuals who receive a CGM product during the observation period. CGM is identified through claims for any CGM related components, including transmitters, sensors, receivers, or readers. This definition constitutes the full sample used in the baseline analysis.

We identify 975 beneficiaries who successfully started CGM during the study period. Rather than modeling CGM adoption, the empirical analysis focuses on continuation of use following the first quarter in which a beneficiary receives a CGM product. For each individual, event time is defined relative to the first observed CGM claim. Quarter 0 corresponds to the quarter of initial initiation, and beneficiaries are followed for up to three additional quarters, yielding a four-quarter (one-year) post-receipt observation window. By definition, all individuals in our sample used a CGM in Quarter 0. This design allows us to examine how reauthorization requirements at the time of first use affect sustained CGM use within the first year after treatment begins.

There are two outcomes of interest. The first outcome is CGM continuation, measured using pharmacy and durable medical equipment claims for CGM components, including transmitters, sensors, receivers, and readers associated with approved CGM systems (Dexcom G6 and Freestyle Libre). Outcomes are based on observed claims rather than prescriptions written, thereby capturing successful authorization and dispensing rather than prescribing intent. The second outcome is health, in which we focus on emergency room visits and hospitalizations, both overall and those with a specific mention of T1D in the claim.

[7] The Early and Periodic Screening, Diagnostic and Treatment benefit in Medicaid requires states to cover CGMs for pediatric beneficiaries without added requirements beyond those in the manufacturers label.

Table 1 presents summary statistics for the sample of beneficiaries with T1D who received a CGM during the study period. The average age is 39 years, with a range from 20 to 64. Approximately 62 percent of the sample is female. The racial composition reflects the demographic distribution of the West Virginia Medicaid population: 83.1 percent identify as White, 3.6 percent as Black, 9.6 percent as another race category, and 3.4 percent as Hispanic.

At the time of first CGM receipt, 24.5 percent of beneficiaries were exposed to the six-month reauthorization policy and 31.9 percent to the three-month policy, with the remainder initiating under the original one-year policy. Alternative classification algorithms for T1D yield broadly similar sample proportions.

The event-time indicators describe CGM continuation patterns relative to the first quarter of CGM initiation. By construction, all beneficiaries are observed in quarter 0, and the share remaining in subsequent quarters declines over time, reflecting discontinuation and attrition in sustained use. Baseline clinical severity of complications, measured using claims observed in the first quarter of CGM use, has a mean of 3.03 and ranges from 0 to 13. This measure captures the number of diabetes-related complications at treatment onset, including hospitalizations, emergency department visits, and diagnosis codes indicative of uncontrolled diabetes.[8] The substantial variation in baseline complication severity indicates meaningful heterogeneity in underlying disease burden at the time of first CGM receipt, motivating the subsequent analysis of heterogeneity by complication severity.

To assess robustness to alternative identification rules, we implement two widely used classification algorithms from the literature, proposed by Klompas et al. (2013) and Teltsch et al. (2019). These algorithms rely on different combinations of diagnosis codes, medication patterns, and utilization history to distinguish T1D from type 2 diabetes in administrative data.[9]

3.2 Empirical Strategy

During the study period, West Virginia Medicaid progressively shortened the CGM reauthorization interval for adult beneficiaries with T1D. Beginning in the fourth quarter of 2021, reauthorization changed from annual to every six months, and in the second quarter

[8]Severity is calculated based on diagnoses of retinopathy-, neuropathy-, nephrology-, cerebrovascular-, cardiovascular-, peripheral vascular-, and metabolic-related conditions. The index is associated with higher risk of diabetes-related hospitalization (Young et al., 2008).

[9]The Klompas algorithm classifies T1D based on a combination of diagnosis codes and insulin use, placing emphasis on the absence of oral hypoglycemic agents. The Teltsch algorithm incorporates diagnosis frequency and treatment patterns, requiring consistent insulin use and limited evidence of type 2 diabetes management.

Table 1: Summary Statistics

	N	Mean	SD	Min	Max
Demographic Characteristics					
Age	975	38.55	11.84	20	64
Female	975	0.621	0.486	0	1
White	975	0.831	0.375	0	1
Black	975	0.0359	0.186	0	1
Other Race	975	0.0964	0.295	0	1
Hispanic	975	0.0338	0.181	0	1
Event-Time Indicators					
Quarter 0	975	1	0	1	1
Quarter 1	968	0.883	0.321	0	1
Quarter 2	745	0.836	0.370	0	1
Quarter 3	649	0.794	0.405	0	1
Policy Indicators at First CGM Receipt					
Twelve-Month Regime	975	0.436	0.496	0	1
Six-Month Regime	975	0.245	0.430	0	1
Three-Month Regime	975	0.319	0.466	0	1
Diabetes Classification Indicators					
Klompas	975	0.579	0.494	0	1
Teltsch	975	0.623	0.485	0	1
Baseline Clinical Severity of Complications					
Baseline Severity	975	3.027	2.681	0	13

Notes: Summary statistics are calculated at the individual level. Policy indicators correspond to the reauthorization regime in effect at the time of first CGM receipt. Baseline severity is constructed using claims observed in the first quarter of CGM use and reflects the count of diabetes-related complications and acute health care utilization indicators, ranging from 0 to 13.

of 2022, it was further shortened to every three months. These policy changes increased the frequency with which beneficiaries were required to obtain physician documentation to maintain CGM access.

We utilize variation in the calendar quarter in which beneficiaries initiated CGM to evaluate the impact of these policy changes on the continued use of CGM. Let k_i^0 denote the calendar quarter of CGM initiation for individual i. Event time is defined relative to CGM initiation, where $k = 0$ denotes the initiation quarter and $k = 1, 2,$ and 3 denote subsequent follow-up quarters within the first year after initiation. We construct indicators for whether CGM initiation occurred under the six-month or three-month reauthorization regimes based on the calendar quarter k_i^0. These indicators are fixed at the individual level and capture exposure to different administrative requirements at the time of initiation. We estimate the following event time specification:

$$\begin{aligned} Y_{itzr} = \alpha_z + \omega_r + \sum_{k=1}^{3} \beta_k \mathbf{1}(EventTime_{it} = k) \\ + \lambda_6 SixMonth_i + \lambda_3 ThreeMonth_i \\ + \sum_{k=1}^{3} \delta_k \mathbf{1}(EventTime_{it} = k) \times SixMonth_i \\ + \sum_{k=1}^{3} \gamma_k \mathbf{1}(EventTime_{it} = k) \times ThreeMonth_i \\ + X_{it}'\theta + \varepsilon_{itzr}. \end{aligned} \tag{1}$$

where Y_{itzr} denotes the outcome for individual i who lives in zip code z and region r in time t. The specification includes zip code fixed effects (α_z) and region fixed effects (ω_r), which account for time-invariant geographic differences in access, provider supply, and health care infrastructure. The vector X_{it} includes age, gender, and race as covariates. Standard errors are clustered at the individual level. The interaction terms allow the evolution of CGM continuation after initiation to vary across reauthorization regimes. Consequently, the coefficients δ_k and γ_k measure how continuation in follow-up quarter k differs under the six-month and three-month reauthorization policies, respectively, relative to the annual reauthorization policy.

Identification relies on the assumption that, conditional on geographic fixed effects and observed covariates, cohorts defined by the quarter of CGM initiation would have followed similar continuation trajectories in the absence of changes in reauthorization requirements.

In particular, no other contemporaneous policy changes or cohort-specific shocks should differentially affect CGM continuation across initiation cohorts. Under this assumption, differences in post-initiation continuation patterns can be attributed to changes in administrative reauthorization frequency. We conduct several additional analyses to assess the robustness of the main findings and the sensitivity of the results to alternative sample definitions, comparison groups, and estimation methods.

Because administrative claims data require a definition of T1D, we assess whether the estimated policy effects are sensitive to alternative classification rules. In addition to our primary claims-based definition, we implement alternative algorithms proposed by Klompas et al. (2013) and Teltsch et al. (2019). These approaches differ in their reliance on diagnosis codes, medication claims, and utilization patterns to distinguish type 1 from type 2 diabetes. For each definition, the sample is restricted accordingly, and the dynamic event-time specification is re-estimated. Consistency between classification rules would indicate that the results are not driven by disease misclassification.

We next isolate the transition to the six-month reauthorization regime. To do so, the sample is restricted to beneficiaries whose first CGM receipt occurred between 2021Q1 and 2021Q4, thereby excluding exposure to the subsequent three-month reform introduced in 2022Q2 and removing beneficiaries whose first CGM use occurred before the COVID-19 pandemic. This narrower policy window provides a cleaner comparison between the annual and six-month reauthorization policies removing aspects of reauthorization that may be related to time trends in CGM initiation. The specification focuses exclusively on the interaction between event time and the six-month policy indicator.

To further distinguish policy effects from general time trends, we compare beneficiaries who first received a CGM under the six-month reauthorization policy with beneficiaries who were already using CGM before the policy change. Existing users are observed over the same calendar period as newly exposed beneficiaries, allowing a contemporaneous comparison within a common time frame. Differences in post-initiation trajectories between these groups provide additional evidence on whether shorter reauthorization intervals reduce sustained CGM use.

We also examine heterogeneity by baseline clinical severity of complications. We construct a severity measure using claims observed during the quarter of CGM initiation, capturing diabetes-related complications and health care utilization at treatment onset. The main specification is estimated across sub-samples defined by alternative severity thresholds, allowing us to assess whether the effects of administrative burden vary with baseline severity (Young et al., 2008).

We also implement a placebo test to assess the validity of the identification strategy. A fictitious reform indicator is assigned to cohorts initiating CGM between 2021Q2 and 2021Q3, when the annual reauthorization policy remained in effect. The sample is restricted to observations before 2021Q4, and the true policy indicators are replaced with this placebo measure. In the absence of an actual policy change, the interaction coefficients should be statistically indistinguishable from zero. Failure to detect a placebo effect supports the interpretation that the main findings are not driven by pre-existing differences in continuation trends across initiation cohorts.

As a final robustness check, we estimate a discrete-time survival model of CGM discontinuation using the same event-time specification as Equation (1), replacing the linear probability outcome with the conditional hazard of first CGM discontinuation. Because discontinuation is observed only at quarterly follow-up intervals rather than continuously, a discrete-time hazard model is more appropriate than a continuous-time Cox proportional hazards model, which assumes continuously observed event times. We therefore estimate a complementary log-log model, the discrete-time analogue of the Cox proportional hazards model, and report hazard ratios. We also estimate an analogous discrete-time logistic model and report the corresponding odds ratios as an additional robustness check.

The survival analysis is restricted to the first observed CGM discontinuation. Beneficiaries remain in the risk set until discontinuation or the end of follow-up. Because the specification includes interactions between event time and reauthorization policy, we report post-estimation hazard ratios and odds ratios obtained from linear combinations of the estimated coefficients, which represent the policy effect at each follow-up quarter relative to the annual reauthorization policy.

Lastly, to assess whether reduced CGM continuation from more frequent reauthorization policies translated into worse health outcomes, we re-estimate equation 1 using four acute healthcare utilization outcomes: all-cause inpatient hospitalizations, diabetes-related inpatient hospitalizations, all-cause emergency department visits, and diabetes-related emergency department visits.

4 Results

4.1 Impact on CGM Continuation

Table 2 presents baseline results estimating equation 1. Column (1) reports results using the full sample, while columns (2) and (3) provide estimates using sub-samples derived from T1D identifying algorithms. The table presents baseline dynamic event-time

estimates of CGM continuation within the first year following the first quarter in which a WV Medicaid beneficiary receives a CGM. Across all specifications, continuation declines over the span of a year. Relative to the first quarter of CGM initiation, the probability of continued CGM use falls steadily in quarters one, two, and three. This pattern reflects natural attrition over time, independent of policy changes.

The key coefficients of interest are the interactions between event time and the reauthorization regimes. These estimates capture whether shortening the administrative renewal window alters the post-treatment trajectory of CGM use. Under the six-month reauthorization policy, continuation declines more sharply beginning in the second quarter after the first CGM claim. Across all three specifications, the interaction term for two quarters post CGM receipt is negative and statistically significant, and the effect becomes larger by the third quarter. The effect size represents a 14 percentage point decrease in CGM continuation in the second quarter. This effect grows to 21 percentage points in quarter three. These represent 17 and 26% reductions in CGM use, respectively. The timing of these effects closely matches the six-month renewal threshold, suggesting that more frequent documentation requirements reduce sustained device use.

The three-month reauthorization policy produces an earlier decline in continuation. The interaction for one quarter is large and statistically significant across all specifications, indicating that discontinuation increases soon after the first required renewal. The second-quarter interaction also remains negative and statistically significant and is of a larger magnitude than that observed under the six-month policy. Because the sample ends in the fourth quarter of 2022, few beneficiaries remain exposed long enough to estimate the third-quarter interaction under the three-month regime, and that term is therefore omitted.

Table 2: Effect of Reauthorization Frequency on CGM Continuation

	(1) Full Sample	(2) Klompas	(3) Teltsch
Event Time (Baseline = Quarter 0)			
1 Quarter	-0.045***	-0.047***	-0.048***
	(0.011)	(0.013)	(0.013)
2 Quarters	-0.086***	-0.083***	-0.085***
	(0.014)	(0.018)	(0.017)
3 Quarters	-0.127***	-0.113***	-0.120***
	(0.017)	(0.020)	(0.020)
Event Time × Six Month			
1 Quarter	-0.035	-0.048	-0.049*
	(0.022)	(0.030)	(0.029)
2 Quarters	-0.139***	-0.123***	-0.122***
	(0.032)	(0.041)	(0.040)
3 Quarters	-0.210***	-0.223***	-0.222***
	(0.037)	(0.048)	(0.047)
Event Time × Three Month			
1 Quarter	-0.199***	-0.214***	-0.212***
	(0.028)	(0.042)	(0.040)
2 Quarters	-0.276***	-0.238***	-0.246***
	(0.054)	(0.071)	(0.068)
Observations	3,328	1,989	2,130
R-squared	0.265	0.292	0.289

Notes: Column (1) reports estimates for the full analytic sample. Columns (2) and (3) apply alternative T1D classification algorithms from the literature. All specifications include zip code and region fixed effects as well as demographic controls (age, gender, and race). Robust standard errors are clustered at the individual level. The interaction term for three quarters under the three month reauthorization regime is omitted due to lack of observations. *** p<0.01, ** p<0.05, * p<0.1.

Our findings indicate that shortening the reauthorization interval reduces sustained CGM use within the first year following CGM initiation. Importantly, the estimated declines emerge immediately after beneficiaries become subject to the corresponding renewal requirement, with no evidence of earlier effects. This close correspondence between policy timing and treatment discontinuation is consistent with administrative burden and ordeal mechanisms acting as a barrier to continued access. These results are consistent across alternative claims-based definitions of T1D, suggesting that the findings are not driven by misclassification of diabetes type.

Table 3 isolates the six-month reauthorization reform to estimate its effect independently of the subsequent three-month policy. Restricting the analysis to this narrower policy window produces estimates that closely mirror the baseline results. Column (1) restricts the sample to cohorts surrounding the implementation of the six-month reform and excludes any exposure to the three-month regime. As in the main specification, continuation declines over event time relative to the first quarter of CGM receipt. The interaction terms indicate that the six-month reform leads to a statistically significant reduction in CGM continuation beginning in the second quarter after receipt, with the effect growing larger by the third quarter. The timing of these effects closely aligns with the six-month renewal requirement, reinforcing the interpretation that increased administrative frequency reduces sustained device use.

Columns (2) and (3) implement a "new" versus "established" CGM patient comparison. These specifications contrast beneficiaries who first received a CGM under the six-month regime with those already using CGM prior to the reauthorization reform but observed over the same calendar window. The negative interaction coefficients remain statistically significant, particularly in the second and third quarters. The results are similar when the sample is further restricted to cohorts beginning in the first quarter of 2021 and later, ensuring comparable pre-reform exposure.

Across all specifications, the estimated decline in continuation begins immediately after beneficiaries become subject to the first six-month renewal requirement. The consistency of the results across alternative comparison groups suggests that the observed effects are unlikely to reflect compositional differences or broader time trends in CGM utilization and instead arise from the increased ordeal mechanism imposed by more frequent reauthorization.

Table 3: Isolating the Six-Month Reauthorization Reform

	(1) Six-Month Reform Window	(2) New vs. Established	(3) New vs. Established (Restricted Cohorts)
Event Time (Baseline = Quarter 0)			
1 Quarter	-0.044**	-0.084***	-0.063**
	(0.021)	(0.016)	(0.025)
2 Quarters	-0.070**	-0.121***	-0.081***
	(0.026)	(0.018)	(0.028)
3 Quarters	-0.140***	-0.204***	-0.235***
	(0.036)	(0.023)	(0.044)
Event Time × Six Month			
1 Quarter	-0.037	0.004	-0.017
	(0.028)	(0.024)	(0.032)
2 Quarters	-0.154***	-0.103***	-0.143***
	(0.039)	(0.034)	(0.041)
3 Quarters	-0.199***	-0.133***	-0.104*
	(0.049)	(0.040)	(0.055)
Observations	1,398	2,359	1,386
R-squared	0.334	0.269	0.331

Notes: Column (1) restricts the sample to initiation cohorts surrounding the implementation of the six-month reauthorization policy and excludes observations exposed to the subsequent three-month reform; the analysis window spans 2021-Q1 through 2022-Q2. Column (2) compares beneficiaries initiating CGM under the six-month regime to beneficiaries already using CGM prior to the reform, observed over the same calendar window. Column (3) further restricts this comparison to initiation cohorts beginning in 2021-Q1 and later to ensure comparable pre-reform exposure. All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors are clustered at the individual level. *** p<0.01, ** p<0.05, * p<0.1.

In Table A1 we run the same three analyses but with the Klompas sample. Using the six-month reform window leads to similar results as before: there is a significant increase in discontinuation for the six-month policy that starts two quarters after first CGM receipt. Comparing “new" versus “established" users in Column (2) now results in smaller effects than before and they are no longer significant, although a 6.6 percentage point decrease in continuation may still be clinically meaningful. When we further restrict the “new" versus “old" comparison, which is shown in Column (3), we still find large and significant declines in CGM continuation (12 percentage points) two quarters after first CGM.

Table A2 reports the results of a placebo test in which a fictitious reform is assigned to cohorts prior to the implementation of the six-month reauthorization policy. As expected, continuation declines over event time relative to the first quarter of CGM receipt. However, the interaction terms associated with the placebo reform coefficients are small and statistically insignificant across all post-event quarters. There is no evidence of differential continuation patterns around the artificial policy threshold.

These findings support our main identification strategy. If the main results were driven by underlying time trends or unobserved cohort differences, we would expect to observe similar effects when assigning a false reform date. The absence of such effects strengthens the interpretation that the observed declines in continuation are attributable to the actual changes in reauthorization frequency rather than spurious temporal dynamics.

We also redo the analyses shown in Table 3, Table A1, and Table A2, but we adjust the margin at which we cluster standard errors. Instead of clustering at the individual level, we now cluster at the zip code level, as patients within the same zip code likely have similar access to medical care. The results are shown in Table A3 and are consistent with more frequent reauthorization leading to reductions in the continuation of CGM. The placebo test still finds no impacts of the fictitious policy change.

Table 4 reports estimates from the discrete-time survival models of first CGM discontinuation. Relative to the annual reauthorization policy, beneficiaries initiating CGM under the six-month reauthorization policy experience more than a threefold increase in the hazard of discontinuation by the second follow-up quarter, coinciding with the timing of the first required reauthorization. Under the three-month policy, the increase occurs one quarter earlier, with beneficiaries experiencing a sevenfold higher hazard of discontinuation, consistent with the shorter reauthorization interval. The discrete-time logistic model yields nearly identical conclusions, indicating that the findings are not sensitive to the choice of discrete-time estimator. Overall, the survival analysis reinforces the baseline results by showing that more frequent reauthorization requirements increase the risk of treatment discontinuation immediately after beneficiaries become subject to the corresponding renewal threshold.

Table 4: Discrete-Time Survival Analysis of CGM Discontinuation

	Complementary Log-Log		Logit	
Policy/Quarter	HR	95% CI	OR	95% CI
Six-Month Reauthorization				
Quarter 1	2.19	[0.90, 4.94]	2.13	[0.91, 5.01]
Quarter 2	3.87***	[2.13, 7.03]	4.25***	[2.24, 8.06]
Three-Month Reauthorization				
Quarter 1	7.00***	[3.58, 13.71]	7.85***	[3.98, 15.50]
Observations	1,609		1,609	

Notes: Entries report hazard ratios (HR) from the complementary log-log model and odds ratios (OR) from the discrete-time logit model. Estimates are obtained from post-estimation linear combinations of the policy and event-time interaction coefficients, and therefore represent the combined effect of each reauthorization policy relative to the annual reauthorization policy at the corresponding follow-up quarter. Only policy-quarter combinations identified in the estimation sample are reported. All specifications include zip code and region fixed effects, demographic controls (age, gender, race, and ethnicity), and robust standard errors clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

We also conduct a robustness check of the discrete-time survival analysis using the Klompas sample, as shown in Table A4. Consistent with the results in the full sample, there are still significant impacts on survival two quarters later for the six-month policy and one quarter after for the three-month policy. Thus, our previous results do not seem to be driven by the method used to determine who has T1D.

Table A5 examines whether the estimated policy effects vary by baseline clinical severity of diabetes-related complications at the time of the first CGM receipt. Across severity subsamples, the overall pattern of declining continuation remains consistent with the main results, indicating this ordeal affects patients across the diabetes complication severity distribution. There is no clear evidence of a monotonic relationship between baseline diabetes-related complication severity and the size of the policy effect.[10] Overall, these findings provide evidence that the adverse effects of more frequent reauthorization requirements are broadly present across patients with differing levels of diabetes-related complications. While previous literature finds that individuals with greater clinical complexity may be particularly affected by administrative burdens (Bell et al., 2023; Christensen et al., 2020),

[10]The zero-severity cohort should be interpreted with caution because of its small sample size; it represents approximately 20% of the analytic sample.

our results show that among Medicaid beneficiaries, patients are impacted by administrative burdens regardless of clinical complexity. Our results are consistent with Wu and Meyer (forthcoming), in which renewal processes do not seem to effectively target those who would benefit least from continuation.

4.2 Impact on Health: Probability of Hospitalization or ER visit

We estimate the effects of these different reauthorization regimes on health outcomes. We first focus on overall hospitalizations and diabetes-related hospitalizations in Table 5. We find a statistically significant 10 percentage points increase in the probability of hospitalization by quarter 3 for those subject to the six-month reform. Point estimates for diabetes-related hospitalizations are slightly diminished, but similarly statistically significant. While this pattern does not exactly match the continuation results, the slightly delayed increase in hospitalization is consistent with worse glycemic control following reauthorization requirements.

We also perform an event study analysis in Table A6, which uses the period immediately before a beneficiary's first receipt of a CGM as the reference period. These results are consistent with those in Table 5 and do not show evidence of pre-trends. While previous research has used variation in prices to show that discontinuing medical treatment leads to increased mortality (Chandra et al., 2024), we use an ordeal mechanism to show that increased treatment discontinuation leads to a higher probability of hospitalization. We find no effect on emergency room visits due to the six-month reform (see Table A7).

We also test for heterogeneity in the effect on hospitalizations by the baseline complication severity, which is shown in Table A8. The result for the full sample, that there is a significant increase in hospitalizations under the six-month reform three quarters after first CGM claim, holds across most of the severity sub-samples. In the cases where the results are no longer significant, it is not because the point estimate is zero, but rather because the standard errors are larger in the smaller sub-samples. For example, in the group with more than four as their severity value, the effect is an 8.4 percentage point increase in hospitalizations, which could still represent a meaningful increase in Medicaid costs. Consistent with the full sample results, when we focus on diabetes-related hospitalizations, the results across severity levels suggest increases, but the magnitudes are smaller in value. Given that part of the reason people can have higher severity levels is hospitalizations, at first glance it may seem like the effect on hospitalizations should be larger among those with the highest severity level. However, there are significant increases in hospitalizations under the six-month reform three quarters after first CGM claim even among those with lower baseline severity. When we focus on diabetes-related hospitalizations, the increase in

Table 5: Effect of Reauthorization Frequency on the Probability of an Inpatient Hospitalization

	All Hospitalizations			Diabetes-Related Hospitalizations		
	Full	Klompas	Teltsch	Full	Klompas	Teltsch
Event Time (*Baseline = Quarter 0*)						
1 Quarter	-0.020	-0.024	-0.020	0.005	0.005	0.008
	(0.020)	(0.025)	(0.024)	(0.018)	(0.024)	(0.023)
2 Quarters	-0.030	-0.040	-0.038	-0.003	-0.015	-0.014
	(0.020)	(0.026)	(0.024)	(0.018)	(0.024)	(0.022)
3 Quarters	-0.074***	-0.082***	-0.078***	-0.034**	-0.047**	-0.045**
	(0.018)	(0.024)	(0.023)	(0.016)	(0.021)	(0.020)
Event Time × Six Month						
1 Quarter	-0.002	-0.015	-0.017	-0.027	-0.043	-0.044
	(0.029)	(0.041)	(0.039)	(0.027)	(0.038)	(0.036)
2 Quarters	-0.007	-0.033	-0.030	-0.051	-0.038	-0.048
	(0.029)	(0.040)	(0.038)	(0.027)	(0.038)	(0.036)
3 Quarters	0.095***	0.095**	0.090**	0.056**	0.067*	0.063*
	(0.028)	(0.038)	(0.036)	(0.026)	(0.038)	(0.035)
Event Time × Three Month						
1 Quarter	-0.013	-0.041	-0.039	-0.035	-0.057	-0.060
	(0.029)	(0.043)	(0.041)	(0.027)	(0.040)	(0.037)
2 Quarters	-0.025	-0.005	-0.006	-0.041	-0.010	-0.015
	(0.040)	(0.067)	(0.061)	(0.038)	(0.064)	(0.058)
Observations	3,774	2,160	2,301	3,774	2,160	2,301
R-squared	0.118	0.156	0.149	0.115	0.145	0.137

Notes: Robust standard errors clustered at the individual level are reported in parentheses. All specifications include age, sex, race, zip code fixed effects, and region fixed effects. The Klompas and Teltsch columns use alternative definitions of Type 1 diabetes. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

magnitude among those with zero complications (12.7 percentage point increase) is larger in magnitude than those with the highest severity level (2.7 percentage point increase), although the 95% confidence intervals overlap. This indicates that it is not only those with the most severe baseline complications who have negative health outcomes under the six-month policy.

Our results indicate that the frequency of reauthorization, not merely the existence of prior authorization, is an independent and economically meaningful driver of treatment discontinuation, and that the timing of discontinuation tracks renewal deadlines closely. This suggests a causal effect of ordeal mechanisms on treatment continuation. Because our setting involves a medical device with a short product cycle and an established, clinically endorsed therapy, these findings suggest that even patients with demonstrated successful use can lose access purely through compliance burden. This implies that states face a tradeoff in terms of short-term fiscal savings versus continuity in care for a treatment that can help keep people out of the hospital.

5 Discussion

Our study shows that shortening Medicaid reauthorization intervals reduces sustained CGM use among adults with T1D. Under the original one-year rule, continuation declines gradually over the first year following the first CGM receipt. When the renewal requirement is shortened to six months, continuation falls more sharply beginning in the second and third quarters. The decline is even larger under the three-month reauthorization regime. The timing of these effects closely matches the renewal thresholds, suggesting that more frequent documentation requirements reduce continued device use.

Our results support more recent literature on ordeal mechanisms that find they may, in fact, decrease targeting efficiency (Deshpande and Li, 2019; Wu and Meyer, forthcoming). Wu and Meyer (forthcoming) importantly point out that while initial certification rules may provide particularly useful ordeals that screen out low value people from using a program, ordeals that require reauthorization may, in fact, reduce targeting efficiency exactly because they are targeting individuals already determined to be high value users through the initial certification process. Our work thus represents a test of their model predictions in a new context. Based on their model, a reauthorization or an increase in the rate of reauthorization required to continue using CGM acts as an increase in cost, leading to an increase in the exit rate from the program (here, an increase in non-continuation of CGM) (Wu and Meyer, forthcoming). This is even though the medical literature has documented that CGM is a cost-effective treatment for T1D (Grunberger et al., 2021; Roze et al., 2020; Wan et al., 2018).

Our results provide support for the Wu and Meyer (forthcoming) model that increasing administrative burdens through a reauthorization ordeal reduces targeting efficiency. There are several reasons for this. First, the Medicaid beneficiaries we study, and their physicians, have already justified the medical necessity of CGM through prior authorization, validating that they are high value patients. Second, shortening reauthorization intervals reduces follow up use, which on its own could be evidence of targeting efficiency. The fact that hospitalization rates increase further provides evidence consistent with targeting inefficiency.

These findings are consistent with the administrative burden literature. Prior work shows that learning and compliance costs can reduce participation in public programs even when eligibility remains unchanged (Arbogast et al., 2024; Herd and Moynihan, 2025). In this setting, beneficiaries must work with physicians to continually compile and submit documentation to maintain CGM access. Because authorization processes are mediated through providers, the results are also consistent with evidence that administrative complexity can

affect access through provider capacity and engagement (American Medical Association, 2025; Dunn et al., 2024; Gaulke et al., 2023).

Beyond administrative burdens, T1D patients face a shortage of endocrinologists, including unequal access based on rurality (Denham et al., 2010; Lu et al., 2015; Romeo et al., 2020; Vigersky et al., 2014). We study a selected group of individuals who already completed the first challenge of using CGM. Previous work demonstrates that patients seeing an endocrinologist are more likely to receive a CGM product (Gaulke et al., 2023). Then, to continue use of CGM, the physician in charge must confirm a patient has been using the CGM correctly, requiring a follow up visit and technology that can work with the CGM system to properly access data.[11] This represents an access to care issue separate from the ordeal mechanism/administrative burden itself, but an issue that interacts with the ordeal to exacerbate reauthorization in a way that also does not differentiate between high and low value patients in a meaningful way to increase targeting efficiency.

Given that a stated policy goal of prior authorization and reauthorization is cost containment, we use cost information from the claims data to determine whether and to what extent we can detect a change in costs to Medicaid. The results are shown in Table A9. The first two columns show the impact on overall Medicaid spending, minus costs on CGM products, as we look at that specifically in other columns.[12] In both the full and Klompas samples, there is no significant impact on overall costs minus CGM after the first CGM claim, but the point estimates are positive and potentially economically meaningful, despite not being precisely estimated. We next break out costs to focus specifically on hospitalization related costs, given we find significant increases in the probability of hospitalization three quarters after the first CGM under the six-month reform. Consistent with the previous result on hospitalization, we find significant increases in costs due to hospitalization (both all cause as shown in columns 3 and 4, and diabetes-specific as shown in columns 5 and 6) three periods after the first CGM claim. Lastly, we estimate the impact on CGM costs in columns 7 and 8. Consistent with our results of decreased CGM usage under the six-month reform, we find CGM costs are statistically significant and negative after the first CGM receipt. Therefore, the effects on CGM continuation and hospitalization are economically meaningful in that we can detect changes in Medicaid spending on those two margins.

[11] For the six-month regime this meant documenting CGM was used daily (twice a day daily for a minimum of 24 days per month) and compliance with patient's "current antidiabetic regimen." Reauthorization was granted for 12 months. For the three-month regime, patients must also show that they had received CGM products through the prior approval duration. They were only granted a six-month continuation.

[12] We winsorize all cost data at the 1st and 99th percentiles.

We next conduct a rough back-of-the-envelope calculation to determine whether this leads to higher costs given that the medical literature concludes CGM is cost effective (Grunberger et al., 2021; Roze et al., 2020; Wan et al., 2018). Using the point estimates from column 1 and column 7 of Table A9 and a standard 3% discount rate employed in the literature, we find an overall increase in the cost to Medicaid. Given that the cost savings occur faster than the increased costs due to hospitalization, an insurer would need to be extremely myopic in terms of their discount rate to avoid this conclusion. In addition to monetary costs, there are likely to be higher administrative costs to an insurer from having to review and process so many additional requests for CGM reauthorization, which should be incorporated into a cost-benefit analysis. An important caveat is that the total costs minus CGM are very imprecisely estimated and are thus not statically significant. However, even if there is no impact on overall costs, the social costs may be enough to overturn the conclusion of cost savings. A formal cost-benefit analysis would account for the reduction in quality adjusted life years given that the health impacts associated with losing access to CGM are severe enough to increase the risk of hospitalization. Even among those retaining CGM access, administrative burdens impose learning, compliance, and psychological costs (Herd and Moynihan, 2018). Doctors also report additional financial costs from prior authorization policies (American Medical Association, 2025). Accounting for the opportunity cost of time spent on additional paperwork would further question whether reauthorization policies truly reduce overall costs. The extent to which these policies save money depends on whose costs one focuses on (Medicaid only versus society more broadly) and the discount rate used across quarters.

Our setting is most similar to Homonoff and Somerville (2021) who study the recertification of SNAP, which is similarly a high-need, more attached sample. While reauthorization policies may help save money on treatment, they also appear to introduce measurable barriers to continued access to a clinically beneficial technology, hurting not only the use of best clinical guidelines but also the health outcomes of patients. This results in a tradeoff for WVM. The results in Table A9 show that WVM saves money from the discontinuation of CGM (last two columns); however, WVM also spends more on hospitalizations and diabetes-related hospitalizations (columns 3-6).

Several limitations should be noted. The analysis relies on claims data and cannot directly observe patient preferences or clinical decisions that may independently influence discontinuation. The study is also based on a single state Medicaid program, and institutional features specific to West Virginia may limit generalizability. Finally, the shorter follow-up window for the three-month regime requires cautious interpretation of longer-term effects. Despite these caveats, the placebo test and multiple robustness checks sup-

port the interpretation that changes in reauthorization frequency contribute to the observed declines in continuation.

Additionally, our study occurs during a period with very little churn as the Families First Coronavirus Response Act (FFCRA) was in place from March 2020 through April 2023. This reduced administrative burden with regard to recertification of Medicaid eligibility. This is a strength in that it reduces competing risks of disenrollment for administrative reasons. However, it also means our estimates are a lower bound on the true effect of administrative burdens overall on continuation of CGM. Previous work demonstrates that these burdens reduce recertification across many programs, including Medicaid and among individuals with chronic disease (Myerson et al., 2025; Wu and Meyer, forthcoming).

6 Conclusion

This paper examines how changes in Medicaid reauthorization requirements affect the continued use of CGM among adults with T1D in West Virginia. Using administrative claims data from 2017 through 2022, we exploit variation in reauthorization rules in the calendar quarter of first CGM usage to estimate dynamic event-time effects under three reauthorization regimes: annual, six-month, and three-month renewal. By focusing on continuation within the first year after initial receipt, the analysis isolates how administrative renewal frequency shapes sustained access to a high-value medical technology. Multiple robustness checks, including alternative diabetes classifications, restricted policy windows, comparisons of new versus established users, and a placebo test, support the identification strategy.

The results show that shortening the reauthorization interval reduces sustained CGM use and increases hospitalization. Under the original one-year policy, CGM continuation declines gradually over time. When the renewal requirement is reduced to six months, continuation falls more sharply, beginning in the second and third quarters following the first CGM receipt. The decline is substantially larger under the three-month reauthorization regime, with pronounced reductions shortly after renewal thresholds are reached. These effects are consistent across alternative sample definitions and are not driven by baseline clinical severity of complications or underlying time trends. The timing of the discontinuation closely aligns with renewal requirements, suggesting that increased administrative frequency contributes to reduced continuation. We also document significant increases in hospitalization under the six-month policy, with the increase starting three quarters after first CGM receipt. We find no evidence of an impact on the probability of an ER visit.

These findings have direct policy implications. They provide evidence that ordeal mech-

anisms decrease benefit use not just on the extensive margin in terms of program take up, but also on the intensive margin of benefit continuation. While reauthorization policies may aim to contain costs or regulate utilization, we provide evidence that increasing the reauthorization frequency increases treatment discontinuation and costly hospitalizations. In publicly insured populations, where administrative processes often intersect with provider capacity constraints and documentation requirements, even modest increases in compliance burden can affect access. Policymakers designing coverage rules for high-value technologies must therefore weigh short-term potential fiscal savings due to treatment discontinuation against increases in costly hospitalizations.

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Appendix

Table A1: Effect of the Six-Month Reauthorization Policy: Klompas Sample

	(1) Six-Month Reform Window	(2) New vs. Established	(3) New vs. Established (Restricted Cohorts)
Event Time (*Baseline = Quarter 0*)			
1 Quarter	-0.051*	-0.074***	-0.053*
	(0.027)	(0.018)	(0.028)
2 Quarters	-0.060*	-0.101***	-0.060**
	(0.028)	(0.021)	(0.030)
3 Quarters	-0.084**	-0.176***	-0.218***
	(0.033)	(0.027)	(0.054)
Event Time × Six Month			
1 Quarter	-0.045	-0.021	-0.042
	(0.039)	(0.033)	(0.040)
2 Quarters	-0.120**	-0.066	-0.120**
	(0.047)	(0.042)	(0.048)
3 Quarters	-0.182***	-0.071	-0.048
	(0.057)	(0.052)	(0.071)
Observations	803	1,448	797
R-squared	0.264	0.234	0.267

Notes: All specifications use the Klompas definition of Type 1 diabetes. Column (1) restricts the sample to initiation cohorts surrounding the implementation of the six-month reauthorization policy and excludes observations exposed to the subsequent three-month reform; the analysis window spans 2021Q1 through 2022Q2. Column (2) compares beneficiaries initiating CGM under the six-month regime to beneficiaries already using CGM prior to the reform, observed over the same calendar window. Column (3) further restricts this comparison to initiation cohorts beginning in 2021Q1 and later to ensure comparable pre-reform exposure. All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors are clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A2: Placebo Test Using Pre-Reform Cohorts

	(1) Full Sample	(2) Klompas	(3) Teltsch
Event Time (*Baseline = Quarter 0*)			
1 Quarter	-0.046***	-0.045***	-0.047***
	(0.013)	(0.016)	(0.016)
2 Quarters	-0.092***	-0.085***	-0.090***
	(0.018)	(0.021)	(0.021)
3 Quarters	-0.123***	-0.106***	-0.119***
	(0.020)	(0.024)	(0.024)
Event Time × *Fake Reform*			
1 Quarter	0.002	-0.006	-0.002
	(0.024)	(0.031)	(0.030)
2 Quarters	0.022	0.010	0.017
	(0.031)	(0.039)	(0.037)
3 Quarters	-0.018	-0.021	-0.003
	(0.041)	(0.047)	(0.046)
Observations	1,676	1,114	1,182
R-squared	0.268	0.330	0.311

Notes: The placebo reform is assigned to initiation cohorts between 2021Q2 and 2021Q3, prior to the implementation of the six-month reauthorization policy. The sample is restricted to pre-reform cohorts only. All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors are clustered at the individual level. None of the interaction terms are statistically significant. *** p<0.01, ** p<0.05, * p<0.1.

Table A3: Robustness Checks Using Zip Code–Clustered Standard Errors

	(1) Full Sample	(2) Klompas	(3) Teltsch	(4) Six-Month Reform Window	(5) New vs. Established	(6) Restricted Cohorts	(7) Placebo Reform
Event Time (*Baseline = Quarter 0*)							
1 Quarter	-0.045***	-0.047***	-0.048***	-0.044**	-0.084***	-0.063**	-0.046***
	(0.010)	(0.012)	(0.012)	(0.018)	(0.014)	(0.025)	(0.012)
2 Quarters	-0.086***	-0.083***	-0.085***	-0.070***	-0.121***	-0.081***	-0.092***
	(0.014)	(0.016)	(0.016)	(0.027)	(0.018)	(0.028)	(0.016)
3 Quarters	-0.127***	-0.113***	-0.120***	-0.140***	-0.204***	-0.235***	-0.123***
	(0.015)	(0.018)	(0.018)	(0.033)	(0.020)	(0.039)	(0.017)
Event Time × Six-Month Reform							
1 Quarter	-0.035*	-0.048*	-0.049*	-0.037	0.004	-0.017	
	(0.020)	(0.028)	(0.028)	(0.025)	(0.023)	(0.030)	
2 Quarters	-0.139***	-0.123***	-0.122***	-0.154***	-0.103***	-0.143***	
	(0.029)	(0.039)	(0.036)	(0.037)	(0.032)	(0.038)	
3 Quarters	-0.210***	-0.223***	-0.222***	-0.199***	-0.133***	-0.104**	
	(0.035)	(0.045)	(0.043)	(0.045)	(0.037)	(0.049)	
Event Time × Three-Month Reform							
1 Quarter	-0.199***	-0.214***	-0.212***				
	(0.026)	(0.035)	(0.035)				
2 Quarters	-0.276***	-0.238***	-0.246***				
	(0.056)	(0.078)	(0.071)				
Event Time × Placebo Reform							
1 Quarter							0.002
							(0.022)
2 Quarters							0.022
							(0.030)
3 Quarters							-0.018
							(0.038)
Observations	3,328	1,989	2,130	1,398	2,359	1,386	1,676
R-squared	0.265	0.292	0.289	0.334	0.269	0.331	0.268

Notes: This table reproduces the primary robustness analyses using standard errors clustered at the zip code level rather than at the individual level. Columns (1)–(3) report the baseline specification using alternative Type 1 diabetes definitions. Columns (4)–(6) report the robustness analyses isolating the six-month reauthorization reform. Column (7) presents the placebo reform test. All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors clustered at the zip code level are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A4: Discrete-Time Survival Analysis of CGM Discontinuation: Klompas Sample

	Complementary Log-Log		Logit	
Policy/Quarter	HR	95% CI	OR	95% CI
Six-Month Reauthorization				
Quarter 1	4.04**	[1.27, 12.83]	3.88**	[1.24, 12.18]
Quarter 2	4.12***	[1.67, 10.16]	4.55***	[1.77, 11.70]
Three-Month Reauthorization				
Quarter 1	7.80***	[2.59, 23.49]	9.16***	[3.05, 27.47]
Observations	803		803	

Notes: Results are based on the Klompas definition of Type 1 diabetes. Entries report hazard ratios (HR) from the complementary log-log model and odds ratios (OR) from the discrete-time logit model. Estimates are obtained from post-estimation linear combinations of the policy and event-time interaction coefficients and therefore represent the combined effect of each reauthorization policy relative to the annual reauthorization policy at the corresponding follow-up quarter. All specifications include zip code and region fixed effects, demographic controls (age, gender, race, and ethnicity), and robust standard errors clustered at the individual level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A5: Heterogeneity by Baseline Severity of Complications at CGM Initiation: CGM Continuation

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Event Time (*Baseline = Quarter 0*)							
1 Quarter	-0.038*	-0.048***	-0.041***	-0.043**	-0.044***	-0.053**	-0.067**
	(0.021)	(0.013)	(0.014)	(0.021)	(0.017)	(0.021)	(0.029)
2 Quarter	-0.087***	-0.086***	-0.072***	-0.060**	-0.083***	-0.090***	-0.090***
	(0.031)	(0.017)	(0.019)	(0.025)	(0.022)	(0.027)	(0.033)
3 Quarter	-0.145***	-0.122***	-0.118***	-0.094***	-0.128***	-0.116***	-0.128***
	(0.039)	(0.020)	(0.024)	(0.030)	(0.027)	(0.030)	(0.039)
Event Time × Six-Month Reform							
1 Quarter	-0.005	-0.042	-0.042	-0.064	-0.061*	-0.014	0.003
	(0.040)	(0.025)	(0.031)	(0.045)	(0.036)	(0.038)	(0.046)
2 Quarters	-0.087	-0.151***	-0.159***	-0.207***	-0.163***	-0.135***	-0.185***
	(0.070)	(0.037)	(0.046)	(0.062)	(0.049)	(0.059)	(0.072)
3 Quarters	-0.119	-0.233***	-0.198***	-0.253***	-0.220***	-0.253***	-0.275***
	(0.083)	(0.042)	(0.052)	(0.069)	(0.056)	(0.068)	(0.080)
Event Time × Three-Month Reform							
1 Quarter	-0.151**	-0.207***	-0.208***	-0.241***	-0.233***	-0.179***	-0.226***
	(0.064)	(0.032)	(0.042)	(0.058)	(0.046)	(0.046)	(0.062)
2 Quarters	-0.196	-0.285***	-0.202**	-0.188*	-0.219**	-0.383***	-0.532***
	(0.134)	(0.060)	(0.080)	(0.099)	(0.080)	(0.086)	(0.108)
Observations	721	2,607	1,684	963	1,472	1,134	806
R-squared	0.376	0.289	0.294	0.398	0.348	0.386	0.480
Sample Restriction	Severity $= 0$	Severity $\neq 0$	Severity ≤ 2	$1 \leq$ Severity ≤ 2	$1 \leq$ Severity ≤ 3	Severity > 3	Severity > 4

Notes: Each column re-estimates the baseline event-time specification on sub-samples defined by baseline severity in the quarter of CGM initiation. Severity is measured using claims-based indicators of acute utilization at treatment onset. All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A6: Effect of Reauthorization Frequency on the Probability of a Hospitalization: Event Study

	All Hospitalizations			**Diabetes-Related Hospitalizations**		
	Full	Klompas	Teltsch	Full	Klompas	Teltsch
Event Time × Six-Month Reform						
4 Quarters Before CGM	-0.045	-0.067	-0.065	-0.060**	-0.090**	-0.086**
	(0.031)	(0.043)	(0.041)	(0.030)	(0.042)	(0.040)
3 Quarters Before CGM	-0.027	-0.029	-0.024	-0.038	-0.052	-0.050
	(0.034)	(0.046)	(0.044)	(0.033)	(0.045)	(0.042)
2 Quarters Before CGM	0.001	-0.000	0.002	-0.009	-0.017	-0.017
	(0.034)	(0.046)	(0.043)	(0.032)	(0.044)	(0.042)
1 Quarter Before CGM (Reference Quarter)	-	-	-	-	-	-
Quarter of First CGM	-0.000	-0.016	-0.011	-0.003	-0.018	-0.017
	(0.033)	(0.045)	(0.043)	(0.030)	(0.042)	(0.039)
1 Quarter After CGM	-0.002	-0.031	-0.028	-0.030	-0.062	-0.061
	(0.031)	(0.043)	(0.041)	(0.030)	(0.041)	(0.039)
2 Quarters After CGM	-0.007	-0.049	-0.042	-0.033	-0.070*	-0.065*
	(0.032)	(0.043)	(0.040)	(0.030)	(0.041)	(0.039)
3 Quarters After CGM	0.094***	0.079*	0.079*	0.052*	0.048	0.046
	(0.033)	(0.047)	(0.045)	(0.031)	(0.046)	(0.043)
Event Time × Three-Month Reform						
4 Quarters Before CGM	0.036	-0.003	-0.006	0.020	-0.021	-0.023
	(0.030)	(0.042)	(0.040)	(0.028)	(0.039)	(0.037)
3 Quarters Before CGM	0.050*	0.070	0.066	0.037	0.058	0.051
	(0.029)	(0.044)	(0.041)	(0.028)	(0.041)	(0.038)
2 Quarters Before CGM	0.042	0.059	0.056	0.030	0.053	0.047
	(0.028)	(0.043)	(0.040)	(0.027)	(0.040)	(0.037)
1 Quarter Before CGM (Reference Quarter)						
Quarter of First CGM	0.097***	0.071	0.071*	0.090***	0.076*	0.075*
	(0.031)	(0.046)	(0.043)	(0.028)	(0.041)	(0.038)
1 Quarter After CGM	0.083***	0.030	0.032	0.054**	0.018	0.014
	(0.028)	(0.037)	(0.035)	(0.027)	(0.035)	(0.033)
2 Quarters After CGM	0.085**	0.082	0.076	0.056	0.076	0.066
	(0.037)	(0.057)	(0.052)	(0.035)	(0.055)	(0.050)
Observations	7,860	4,375	4,627	7,860	4,375	4,627
R-squared	0.115	0.132	0.129	0.098	0.120	0.114

Notes: Robust standard errors clustered at the individual level are reported in parentheses. All specifications include age, sex, race, zip code fixed effects, and region fixed effects. The Klompas and Teltsch columns use alternative definitions of Type 1 diabetes. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A7: Effect of Reauthorization Frequency on the Probability of an Emergency Department Visit

	All ER Visits			**Diabetes-Related ER Visits**		
	Full	Klompas	Teltsch	Full	Klompas	Teltsch
Event Time (*Baseline = Quarter 0*)						
1 Quarter	-0.017	-0.031	-0.026	-0.012	-0.044*	-0.042*
	(0.026)	(0.031)	(0.029)	(0.022)	(0.027)	(0.025)
2 Quarters	-0.019	-0.046	-0.040	-0.009	-0.021	-0.020
	(0.025)	(0.029)	(0.028)	(0.021)	(0.027)	(0.026)
3 Quarters	-0.021	-0.039	-0.037	-0.032	-0.049*	-0.046*
	(0.026)	(0.034)	(0.032)	(0.022)	(0.028)	(0.027)
Event Time × Six Month						
1 Quarter	0.024	0.016	0.012	0.047	0.040	0.038
	(0.043)	(0.056)	(0.053)	(0.038)	(0.051)	(0.048)
2 Quarters	-0.001	-0.010	-0.006	-0.027	-0.080	-0.075
	(0.043)	(0.055)	(0.053)	(0.037)	(0.049)	(0.046)
3 Quarters	0.007	0.079	0.074	0.013	0.043	0.041
	(0.043)	(0.058)	(0.055)	(0.035)	(0.048)	(0.045)
Event Time × Three Month						
1 Quarter	-0.071*	-0.073	-0.068	-0.083**	-0.105**	-0.099**
	(0.041)	(0.060)	(0.055)	(0.037)	(0.055)	(0.050)
2 Quarters	-0.128**	-0.146**	-0.127*	-0.103**	-0.150**	-0.125*
	(0.056)	(0.077)	(0.073)	(0.049)	(0.074)	(0.070)
Observations	3,774	2,160	2,301	3,774	2,160	2,301
R-squared	0.151	0.215	0.208	0.144	0.212	0.203

Notes: Robust standard errors clustered at the individual level are reported in parentheses. All specifications include age, sex, race, zip code fixed effects, and region fixed effects. The Klompas and Teltsch columns use alternative definitions of Type 1 diabetes. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A8: Heterogeneity by Baseline Severity of Complications at CGM Initiation: Hospitalization Outcomes

	All Hospitalizations							Diabetes-Related Hospitalizations						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
Event Time (*Baseline = Quarter 0*)														
1 Quarter	-0.056	-0.009	-0.000	0.048	0.041	-0.066*	-0.046	-0.019	0.012	0.030	0.072**	0.062**	-0.044	-0.014
	(0.047)	(0.022)	(0.028)	(0.036)	(0.031)	(0.034)	(0.041)	(0.039)	(0.021)	(0.026)	(0.036)	(0.029)	(0.031)	(0.037)
2 Quarters	-0.055	-0.023	-0.034	-0.016	-0.041	-0.004	0.035	-0.027	0.004	-0.013	0.000	-0.010	0.018	0.066
	(0.045)	(0.023)	(0.026)	(0.031)	(0.027)	(0.040)	(0.048)	(0.038)	(0.021)	(0.022)	(0.028)	(0.025)	(0.036)	(0.042)
3 Quarters	-0.093***	-0.068***	-0.059**	-0.030	-0.054**	-0.081**	-0.058	-0.056*	-0.028	-0.030	-0.006	-0.019	-0.035	-0.001
	(0.035)	(0.021)	(0.023)	(0.033)	(0.027)	(0.034)	(0.044)	(0.029)	(0.019)	(0.019)	(0.026)	(0.022)	(0.031)	(0.038)
Event Time × Six-Month Reform														
1 Quarter	0.035	-0.012	-0.048	-0.113**	-0.066	0.049	0.072	-0.003	-0.034	-0.070**	-0.124***	-0.087**	0.027	0.041
	(0.062)	(0.033)	(0.037)	(0.047)	(0.041)	(0.057)	(0.065)	(0.057)	(0.032)	(0.034)	(0.046)	(0.038)	(0.055)	(0.063)
2 Quarters	0.034	-0.017	-0.006	-0.036	-0.008	-0.024	-0.034	0.006	-0.039	-0.019	-0.039	-0.031	-0.046	-0.065
	(0.061)	(0.033)	(0.036)	(0.047)	(0.041)	(0.055)	(0.064)	(0.056)	(0.031)	(0.033)	(0.043)	(0.039)	(0.053)	(0.060)
3 Quarters	0.164***	0.078**	0.078**	0.017	0.046	0.115**	0.084	0.127**	0.039	0.056*	0.006	0.011	0.070	0.027
	(0.055)	(0.032)	(0.034)	(0.046)	(0.041)	(0.054)	(0.064)	(0.052)	(0.031)	(0.033)	(0.044)	(0.037)	(0.052)	(0.060)
Event Time × Three-Month Reform														
1 Quarter	0.056	-0.029	-0.014	-0.070	-0.094**	0.038	0.015	-0.000	-0.044	-0.037	-0.072*	-0.099***	0.015	-0.016
	(0.063)	(0.033)	(0.037)	(0.047)	(0.042)	(0.051)	(0.061)	(0.054)	(0.031)	(0.031)	(0.040)	(0.038)	(0.049)	(0.058)
2 Quarters	0.004	-0.026	0.036	0.068	0.057	-0.093	-0.197**	-0.047	-0.036	0.030	0.102	0.052	-0.105	-0.216***
	(0.073)	(0.047)	(0.056)	(0.090)	(0.063)	(0.072)	(0.083)	(0.067)	(0.045)	(0.051)	(0.078)	(0.058)	(0.070)	(0.079)
Observations	733	3,041	1,740	1,007	1,561	1,479	1,099	733	3,041	1,740	1,007	1,561	1,479	1,099
R-squared	0.312	0.120	0.211	0.258	0.195	0.187	0.224	0.292	0.116	0.199	0.226	0.173	0.193	0.232
Sample Restriction	Sev.=0	Sev.≠0	Sev.≤2	1 ≤Sev.≤2	1 ≤Sev.≤3	Sev.> 3	Sev.> 4	Sev.=0	Sev.≠0	Sev.≤2	1 ≤Sev.≤2	1 ≤Sev.≤3	Sev.> 3	Sev.> 4

Notes: Each column re-estimates the baseline event-time specification on subsamples defined by baseline severity at CGM initiation. Panels report results for all-cause inpatient hospitalizations (Columns 1–7) and diabetes-related inpatient hospitalizations (Columns 8–14). All specifications include zip code and region fixed effects and demographic controls (age, gender, and race). Robust standard errors clustered at the individual level are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A9: Effect of Reauthorization Frequency on Costs

	All Costs (No CGM)		**Hospitalization Costs**		**Diabetes-Related Hospitalization Costs**		**CGM Costs**	
	Full	Klompas	Full	Klompas	Full	Klompas	Full	Klompas
Event Time × Six-Month Reform								
4 Quarters Before CGM	610.211	291.635	193.815	165.999	221.678	267.368	2.943	-4.648
	(997.771)	(729.228)	(257.085)	(375.144)	(232.243)	(338.204)	(6.153)	(11.524)
3 Quarters Before CGM	1,689.481***	1,961.432**	649.165**	824.317**	495.452*	571.878	6.928	5.223
	(605.343)	(818.167)	(265.070)	(370.173)	(259.066)	(363.947)	(5.130)	(9.762)
2 Quarters Before CGM	628.374	545.330	339.463	315.384	355.488	284.346	-1.411	-9.771
	(542.178)	(651.858)	(290.386)	(377.097)	(259.911)	(340.335)	(3.578)	(7.467)
1 Quarter Before CGM (Reference Quarter)	-	-	-	-	-	-	-	-
Quarter of First CGM	109.094	320.786	152.709	137.444	165.273	140.440	-142.804***	-194.346***
	(524.121)	(739.652)	(241.385)	(334.664)	(230.019)	(320.845)	(44.510)	(58.035)
1 Quarter After CGM	415.257	841.754	343.107	399.473	240.661	233.332	-126.781***	-159.763***
	(656.729)	(854.641)	(239.014)	(342.994)	(213.867)	(303.727)	(42.995)	(55.897)
2 Quarters After CGM	477.875	408.838	283.893	213.103	199.956	108.164	-277.806***	-317.348***
	(720.931)	(893.135)	(279.667)	(395.030)	(252.038)	(354.372)	(45.164)	(57.581)
3 Quarters After CGM	348.990	848.814	711.935***	749.916**	589.989**	630.323*	-198.188***	-222.484***
	(700.243)	(893.777)	(272.445)	(373.768)	(253.770)	(349.840)	(49.859)	(64.240)
Observations	5,009	3,147	5,009	3,147	5,009	3,147	5,009	3,147
R-squared	0.250	0.233	0.102	0.118	0.099	0.116	0.645	0.675

Notes: All columns focus on the comparison with the six-month reform only. The analysis only goes through 2022-Q2. Because cost data are highly skewed, we winsorize all cost measures at the 1st and 99th percentiles. Robust standard errors clustered at the individual level are reported in parentheses. All specifications include age, sex, race, zip code fixed effects, and region fixed effects. The Klompas columns use alternative definitions of Type 1 diabetes. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.