The Impact of Physician Fee Schedule Changes in Workers Compensation: Evidence From 31 States

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Abstract
Motivation. Quantifying the effects of changes to physician fee schedules in workers compensation has become an integral part of NCCI legislative pricing, as an increasing number of jurisdictions have introduced such legal provisions over the past few decades. This study measures the impact of changes to physician fee schedules on the price and utilization levels of medical services consumed in the context of workers compensation.
Method. The effect of changes to the price ceilings imposed by physician fee schedules is quantified using an impulse-response time series framework. The analysis is based on data for 31 jurisdictions and 11 years (2000–2010). For the purpose of the statistical analysis, monthly price, utilization, and severity indexes are developed, from which rates of inflation and rates of utilization and severity increases are computed. When gauging the response, state-specific characteristics are considered using the concept of price departure and, alternatively, the differential in the price level (at reimbursed amounts) relative to neighboring jurisdictions. The price departure measures the percentage difference between the price level at reimbursed amounts and the price level implied by the MAR (maximum allowable reimbursement) specified in the fee schedule.
Results. On average, in response to a fee schedule increase (that is, an increase in MAR), severity increases by about 76 percent of the original impulse (or 81 percent, depending on the statistical model). This impulse equals the product of the percentage fee schedule increase and the proportion of transaction volume subject to the fee schedule. The magnitude of the response is greater (less) than the global value of 76 percent where the price departure prior to the fee schedule change is comparatively narrow (extensive). For fee schedule decreases, the severity response equals about 45 percent of the impulse. The severity adjustment being less than 100 percent of the impulse is largely due to an inelastic price level adjustment, as there is no material and lasting utilization response to fee schedule changes.
Availability. The statistical models were implemented in R (cran.r-project.org), using the sampling platform JAGS (Just Another Gibbs Sampler, mcmc-jags.sourceforge.net). JAGS was linked to R via proprietary software that makes use of parallelization.

Keywords. Medical Price Inflation, Physician Fee Schedules, Severity, Utilization, Workers Compensation

1. INTRODUCTION

Quantifying the effects of changes to physician fee schedules in workers compensation has become an integral part of NCCI legislative pricing, as an increasing number of jurisdictions have introduced such legal provisions over the past few decades. This study measures the impact of changes to physician fee schedules on the price and quantity levels of medical services consumed in the context of workers compensation.

The effect of changes to the price ceilings imposed by physician fee schedules is quantified using an impulse-response time series framework. The analysis comprises 31 jurisdictions and 11 years (2000–2010). For the purpose of the statistical analysis, monthly price, utilization, and severity indexes are developed, from which rates of inflation and rates of utilization and severity increases are computed. When gauging the response, state-specific characteristics are considered using the concept of price departure and, alternatively, the differential in the price level (at reimbursed amounts) relative to neighboring jurisdictions. The price departure measures the percentage difference between the price level at reimbursed amounts and the price level implied by the MAR (maximum allowable reimbursement) laid down in the fee schedule.

On average, in response to a fee schedule increase (that is, an increase in MAR), severity increases by around 76 percent of the original impulse (or 81 percent, depending on the statistical model). This impulse equals the product of the percentage fee schedule increase and the proportion of transaction volume subject to the fee schedule—for the purpose of this study, the term transaction volume refers to the dollar-weighted sum of quantities of service. The magnitude of the response is greater (less) than the global value of 76 percent where the price departure prior to the fee schedule change is comparatively narrow (extensive). For fee schedule decreases, the severity response equals about 45 percent of the impulse. The severity adjustment being less than 100 percent of the impulse is largely due to an inelastic price level adjustment, as there is no material and lasting utilization response to fee schedule changes. (In general, the defining characteristic of an inelastic adjustment is a percentage change in the effect that falls short of the percentage change in the cause in absolute value terms.)

1.1 Research Context

To the best of the authors' knowledge, there is no published research on the effect of workers compensation physician fee schedule changes on the price and quantity levels of medical services consumed by claimants. On the other hand, the effects of introductions of physician fee schedules in workers compensation are well studied. Schmid and Lord [10] offer an overview of this literature, alongside an analysis of their own for Tennessee and Illinois. In their case study, Schmid and Lord are unable to establish evidence of material changes in the level of consumption of medical services in response to the implementation of physician fee schedules—this finding agrees with previous studies, as discussed in Schmid and Lord. The absence of utilization effects may (entirely or in part) be related to cost-containment measures that were introduced alongside the physician fee schedule or were in place at the time the fee schedule became effective.

There are several studies on physician responses to Medicare fee schedule changes using micro-level data (that is, data at the level of the physician)—for a recent analysis, see Hadley et al. [2], and, for a survey of the literature, see Hadley and Rechosky [3]. The evidence obtained from this line of research regarding the supply response of physicians is largely inconclusive. Modern studies that are based on microeconometric models in the tradition of McGuire and Pauly [9] tend to find evidence of a decrease in the supply of medical services billed to Medicare in response to a Medicare fee schedule decrease. This response agrees with classical economics, which maintains that a price decline is met with a decrease in supply. Then again, there are earlier, less elaborate studies, the findings of which support the income-targeting hypothesis. The income-targeting hypothesis posits that physicians, in response to a decrease in the Medicare fee schedule, increase the supply of services billed to Medicare in an attempt to maintain a chosen level of revenue. In some of these earlier studies, by relying on ad hoc regression approaches, the uncovered economic relations may be due to spurious correlation as utilization trends (overall or for individual CPT codes) are misread as responses to fee schedule decreases. For a brief summary of these two strands of literature, see Congressional Budget Office [1].

An important insight delivered by McGuire and Pauly [9] is the interdependence of the markets for medical services when there is more than one potential payer. This interdependence was highlighted in a recent study by Heaton [4] on the effects of the 2006 Massachusetts healthcare reform. Heaton's research shows that the billing of emergency room visits to workers compensation is related to the availability of Medicaid.

The study presented here does not account for interdependence between workers compensation and other payers, such as Group Health, Medicare, or Medicaid. The primary reason for not modeling such interdependence is the lack of micro-level (that is, physician-level) data. Although for some physicians, medical services delivered in the context of workers compensation may amount to a significant proportion of total business, the workers compensation system overall accounts for only 1.3 to 2.1 percent of medical expenditures in the economy during the period analyzed; these percentages are based on annual data published by the Centers for Medicare and Medicaid Services (cms.gov). Therefore, in the absence of micro-level data, there is little prospect of uncovering a spillover of fee schedule changes into other channels of provision of medical services. Then again, for the purpose of quantifying the effect of (workers compensation) fee schedule changes on the price level and the consumption of medical services provided by physicians in the context of workers compensation, disregarding such a potential spillover does not compromise the findings.

The analysis does not make use of a general equilibrium framework. Clearly, in the long run, the elasticity of the price level at reimbursed amounts with respect to the price level at fee schedule must equal unity; put differently, over a long time horizon, the compound average rates of inflation must be equal to each other. This perspective of a stable long-term relation between the two rates of inflation suggests an error-correction model where fee schedule changes are treated as shocks that dissipate over time as the price departure settles back into its long-term equilibrium. Then again, the 11-year time period may be too short to identify such a long-term relation between the price indexes at fee schedule and at reimbursed amounts. As will be demonstrated for Georgia, the price departure can keep increasing for many years without signs of mean-reversion.

Similarly, the study does not specify a reaction function of the policymaker. Yet, and in spite of the observed ever-increasing price departure in Georgia, it can be hypothesized that fee schedule changes are responsive to market conditions and, in particular, to the magnitude of the prevailing price departure or the difference in price levels between the state and its neighbors. If the policymaker is sensitive to current market conditions but no reaction function is specified, an endogeneity bias may result. As a result of this endogeneity bias, the parameters that capture the market response to fee schedule changes are biased toward zero, thus causing the strength of the effect to be underestimated.

1.2 Objective

The objective of this study is to evaluate the impact of changes to physician fee schedules on the medical costs of workers compensation claims. The medical costs of these claims are measured using a newly developed concept of contemporaneous severity. This severity concept reflects the price and the quantity levels of medical services consumed within a given time window, where the consumption is normalized by the number of claims that are active during this time interval. As a result of the contemporaneous nature of this severity measure, changes in the consumption of medical services that arise from variations in claim duration are not accounted for.

The study covers 31 jurisdictions over a time period of 11 years. Even though the analysis accounts for differences across states (albeit in limited ways), the estimated responses represent what the policymaker can expect to materialize on average. The effect of any single historical (or future, out of sample) physician fee schedule change may deviate from the estimated response in significant ways. Only when averaged over repeated evaluations of fee schedule changes can the policymaker expect that the observed effects agree with the parameters estimated in this study.

1.3 Outline

What follows in Section 2 is a description of the data on fee schedules and medical transactions that are employed in this study. Section 3 outlines the price, utilization, and severity indexes used to gauge the impulses of fee schedule changes and responses thereto; further, this section introduces the concept of price departure. Next, Section 4 offers charts with descriptive statistics on price inflation and the rate of utilization increase. Then, Section 5 discusses the statistical models employed in the estimation of the fee schedule responses before presenting the findings. Section 6 offers conclusions. Finally, Section 7 provides an appendix.

2. THE DATA

The study uses data from two sources. First, physician fee schedules were obtained from Ingenix, Inc. (now known as OptumInsight, Inc.) and employed in the analysis as reported by this data provider. Second, the effects of fee schedule changes are measured on a large set of medical transactions associated with workers compensation claims. These transaction records cover the time period January 1, 2000 through December 31, 2010 and were provided by a set of (primarily large) insurance carriers. The first rates of change in price level, utilization, and severity apply to February 2000 (over January 2000).

The effect of fee schedule changes is analyzed for 31 jurisdictions. For five additional jurisdictions, which do not have fee schedules in place during the period analyzed, descriptive statistics is provided. See Section 7.1 in the appendix for a list of the 36 jurisdictions included in this study.

The jurisdiction state criterion and provider zip code information are used when associating observed medical transactions with a given state. As measured by written premium, for the year 2006, the market share of the carriers contributing to the medical transactions data set ranges between 15 percent and 80 percent, depending on the jurisdiction. Carriers contributing data include some state funds, but not all, in the jurisdictions analyzed. Self-insureds are not included in this market share computation.

Three states introduced fee schedules during the period analyzed. For another 14 states, the start date of the analysis predates the first fee schedule considered in the study; for details, see Section 7.1 in the appendix.

The data set excludes transactions associated with medical services provided by hospitals and ambulatory surgical centers, but includes transactions related to services delivered by physicians (as the provider type) at these places of service. The medical transactions data were edited using expert knowledge on billing and reimbursement practices, and the data set was cleansed using statistical tools of outlier detection. For an overview on the data-cleansing tools, see Sections 7.2 through 7.5 in the appendix.

The medical services associated with the obtained transaction records can be categorized into the American Medical Association (AMA) service categories Evaluation and Management Services, Anesthesia, Surgery, Radiology, Pathology and Laboratory, and Medicine. Transactions related to Anesthesia are excluded from the study due to difficulties in quantifying the units of service that are associated with the individual records. The service category Pathology and Laboratory, although sparsely populated, is included.

For the purpose of this study, medical services are identified by a combination of CPT code and modifier; only modifiers that are recognized by fee schedules are considered. For transactions associated with a thus identified medical service, the MAR may vary by geozip (geographic areas identified by arrays of zip codes; Florida, Illinois, and Texas only) and place of service. Where such variation exists, the MAR of a given medical service cannot be read directly from the fee schedule but instead needs to be calculated for any given month as a weighted average across geozips and places of service, where the weights are the number of units of service provided.

For a given medical service, the study recognizes a MAR only if the fee schedule specifies the price ceiling as a dollar amount (dubbed fixed-value MAR), as opposed to setting the reimbursement limits by what is considered usual and customary or by defining it as a percentage of billed charges, for instance. When a fee schedule change occurs mid-month, for the purpose of calculating the average MAR of a given medical service for that month, the pertinent fee schedules are prorated based on the numbers of units of service provided under each regime.

Because the MAR that applies to a given medical transaction may differ by the place of service and (in Florida, Illinois, and Texas) across geozips, the average monthly MAR of a given medical service may vary over time for a given fee schedule. This is because the distribution of transactions by place of service and geozip may change from month to month.

3. RESEARCH FRAMEWORK

The impulse of and the response to fee schedule changes are read from indexes. These indexes are on a monthly basis and calculated for the mentioned AMA categories and for All Categories. Only results for All Categories are presented. Further, the concept of price departure is introduced as a (monthly) measure of deviation between the price level at reimbursed amounts and the price level implied by the price ceilings (that is, MAR) laid down in the fee schedule.

3.1 Price, Utilization, and Severity Indexes

Price indexes are calculated from three sets of prices. First, there is a price index calculated from reimbursed amounts, encompassing all medical services, labeled “All CPT.” Second, there is a price index that comprises only medical services subject to a fixed-value MAR, labeled “Subjected CPT.” Third, there is a price index “at fee schedule,” which is based on the stipulated fixed-value MAR. When no fixed-value MAR is available for a given CPT code (for instance, because this medical service is subject to usual and customary reimbursement or because the price ceiling is defined as a percentage of billed charges), this medical service does not contribute to the computation of the price index at fee schedule (or the price index at reimbursed amounts, subjected CPT). Unless noted otherwise, the price index and inflation concepts refer to the reimbursed amounts, All CPT.

All price indexes are Fisher indexes—prices and quantities are measured on the level of the medical service. Technically, the Fisher index is the geometric mean of the respective Laspeyres and Paasche indexes. The Laspeyres index compares the set of prices of the current month to the set of prices of the previous month, using as weights the quantities of medical services of the previous month. The Paasche index undertakes this comparison by using as weights the quantities of the current month.

The formula for the Laspeyres index reads

$$P_L = \frac{\sum_i p_i q_{i-1}}{\sum_i p_{i-1} q_{i-1}} \ . \tag{1}$$

By comparison, the equation for the Paasche index is given by

$$P_P = \frac{\sum_i p_i q_i}{\sum_i p_{i-1} q_i} \ . \tag{2}$$

In the equations above, swapping prices for quantities and vice versa delivers the Laspeyres and Paasche quantity indexes. Analogous to the Fisher price index, the Fisher quantity index is obtained as the geometric mean of the Laspeyres and Paasche quantity indexes. For details on price and quantity indexes, see International Labour Office [8].

The Fisher index has a host of desirable properties. The most important property of the Fisher index in the context of this study is the ability of this index to accurately break down the price and quantity responses to changes in relative prices. For instance, if the structure of consumed medical services changes in response to some services increasing in price less than others, then the Laspeyres (Paasche) price index overestimates (underestimates) the rate of inflation in the event that medical consumption shifts toward services that have increased in price comparatively less. This bias occurs because the Laspeyres index is based on the quantities of the previous month, which do not yet account for the shift in consumption toward the services that have become comparatively less expensive. Conversely, the Paasche index draws on the new quantities, thereby biasing the reading of the rate of inflation in the opposite direction.

The Fisher quantity index corresponds to the Fisher price index at reimbursed amounts, All CPT. The product of these two Fisher indexes is related to the transaction volume. Specifically, the percentage change from the previous month of the product of the Fisher price and quantity indexes equals the percentage change in the transaction volume on which these indexes are calculated.

A utilization index is calculated by normalizing the Fisher quantity index by the number of active claims in the applicable month. In keeping with the price indexes, the utilization index is calculated for the mentioned service categories and for All Categories. Note that the number of active claims in a given month may vary across service categories. Specifically, a claim is considered active in a given service category (in All Categories) if there is a transaction associated with this claim in this service category (in any service category) that enters the price index for this service category (for All Categories) in that month. In accordance with the price index concept, the number of active claims is calculated using the "month *t* to month *t*-1 ratio."

A severity index is calculated as the product of the Fisher price index at reimbursed amounts and the utilization index. The following relation holds among the rates of severity growth (s), the rate of inflation (p), and the rate of utilization growth (u):

$$s = (1+p)\cdot(1+u) - 1 \ . \tag{3}$$

The utilization index represents a contemporaneous concept of utilization, since it builds only on transactions that have been observed in the month under consideration. To the degree that the consumption of medical services provided by physicians is front-loaded in the lifetime of a claim, the utilization index (and, hence, the severity index) may increase in response to an influx of claims. This property of the utilization index may explain some of the seasonal variation that the utilization index (and, as a result, the severity index) displays over the course of the calendar year. Seasonality in the utilization of medical services is to be expected, given the influence of climactic conditions on economic activity, in particular in the construction and leisure and hospitality industries. Finally, to the degree that there is front-loading of medical services in the lifetime of a claim, the utilization index may decrease in response to a systematic increase in claim duration.

Minor simplifications apply. In Tennessee, for instance, in Physical Medicine, the MAR of a given medical service may be regressive with respect to the number of units of service provided to the claimant. Due to data limitations, this regression in MAR was not factored into the computation of the price index at fee schedule.

3.2 Price Departure

Typically, medical services provided by physicians are reimbursed at or below fee schedule, although jurisdictions vary by the degree to which the stipulated price ceilings are enforced. For instance, in Tennessee, the Commissioner may assess civil penalties for fee schedule violations at his discretion. By comparison, in Illinois, reimbursement above fee schedule is permitted when agreed to pursuant to a written contract.

Price departure measures the deviation of reimbursement from fee schedule. Specifically, price departure is defined as the ratio of transaction volume at reimbursed amounts to transaction volume at fee schedule, minus 1. To provide a numerical example, a price departure of a negative 0.05 states that, on average, the reimbursed amounts are 5 percent below fee schedule. In the denominator, for transactions that are associated with medical services not subject to a fixed-value MAR, reimbursed amounts substitute for the price ceilings. Note that the transaction volume that contributes to the price departure calculation is slightly more comprehensive than the transaction volume that finds its way into the computation of the price and quantity indexes. This is because medical services enter the price (or quantity) index of a given month only if there is at least one transaction associated with these services in both the current month and the prior month.

Changes in the difference between the price indexes at reimbursed amounts and at fee schedule do not map exactly into changes in price departure. For instance, for a given array of prices, the price index does not respond to changes in quantities. By contrast, the price departure may increase for given reimbursement practices and a given fee schedule if the distribution of transactions shifts in favor of medical services that exhibit a comparatively large spread between reimbursed amount and MAR.

In keeping with the computation of the price index at fee schedule, for Tennessee, the regressive nature of the fee schedule in Physical Medicine was not factored into the computation of the price departure.

4. DESCRIPTIVE STATISTICS

Descriptive statistics on price level, utilization, and severity increases are presented for the entire set of 36 jurisdictions. Then, in preparation of the statistical analysis of the 31 jurisdictions with fee schedules, charts describe the trajectories of price level, utilization, severity, and price departure for Florida and Georgia.

4.1 General

In what follows, the rate of growth of the price level (i.e., the rate of inflation) is contrasted with the rate of growth of the Medical Care component of the Consumer Price Index (All Urban Consumers: Medical Care for the U.S. city average) and the rate of utilization growth. Further, the actual rate of inflation is compared to the rate of inflation implied by the (fixed-value) MAR laid down in the fee schedule. The rates of growth displayed in the charts are compound annual growth rates (CAGR). For the CAGR concept, the relation displayed in Equation (3) holds.

Chart 1 compares for each of the 36 jurisdictions the rate of inflation of physician services (All CPT) provided in the context of workers compensation to the Medical Care and Physician Services components of the CPI—these components are taken from the CPI for the United States (as opposed to being based on the applicable regional CPI index series). The first monthly price index value applies to January 2000, and the final value is for December 2010. For jurisdictions that introduced a fee schedule during this time window and for which fee schedule information became available during this time interval, the first price level that enters the CAGR calculation pertains to the third complete month following the fee schedule introduction and availability, respectively. Non-fee schedule states are jurisdictions that do not have fixed-value MAR in place at any time during the period analyzed—the absence of fixed-value MAR for services provided by physicians does not imply that there exists no price regulation for medical services provided by physicians based on the charged amount or what is considered usual and customary, nor does it imply that there exist no fee schedules for hospitals or other nonphysician entities.

Chart 1: Rates of Price Inflation, February 2000–December 2010, CAGR

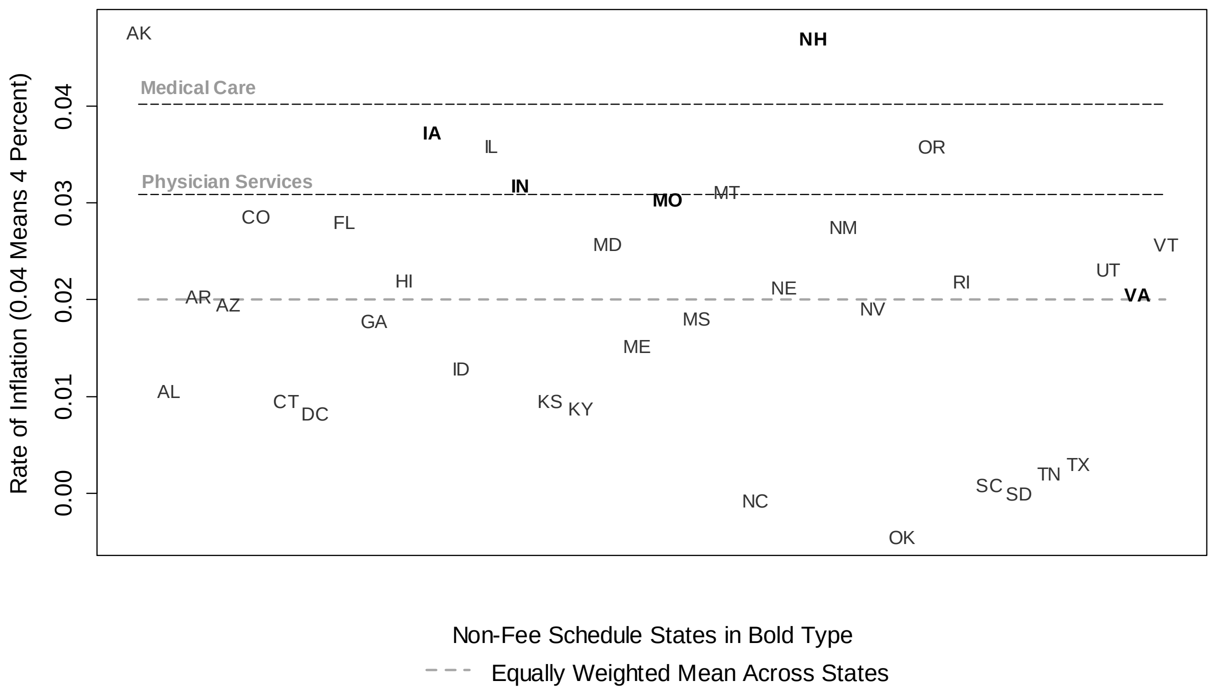
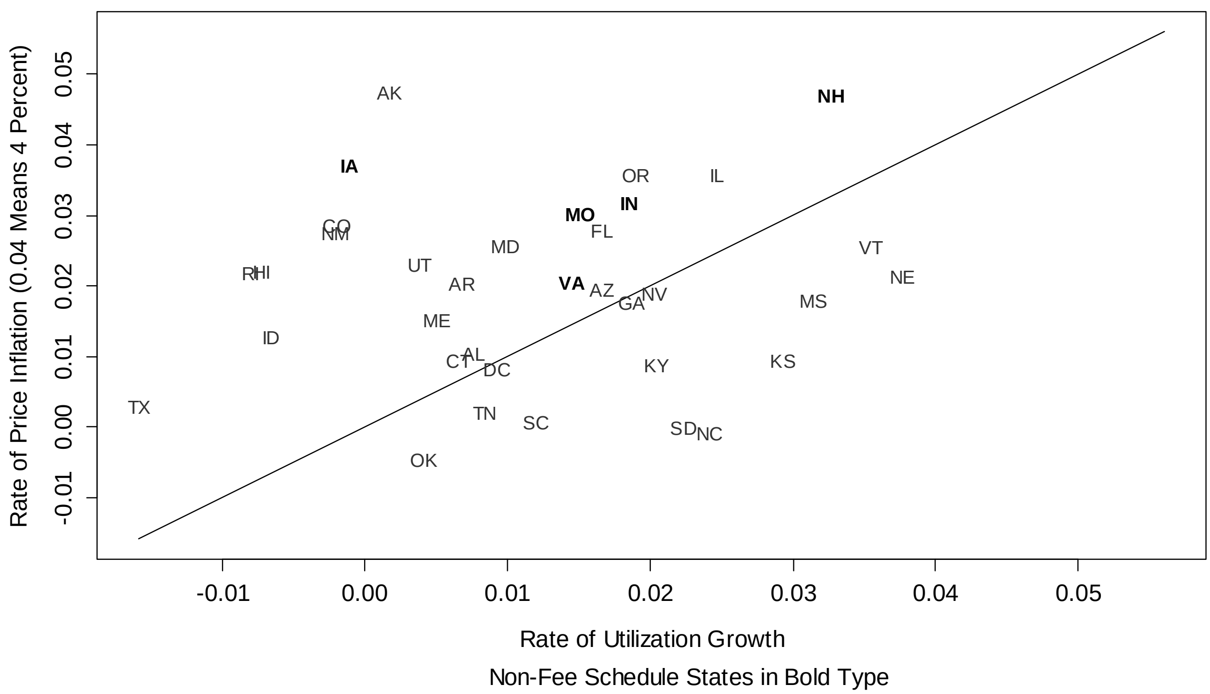


Chart 1 demonstrates that, for many states, the rate of price inflation of physician services in workers compensation is materially below the rate of inflation in the Medical Care and Physician Services components of the CPI. In part, this difference may be due to methodological differences between the CPI and the Fisher index presented here. For methodological details on the role of medical services in the CPI, see bls.gov/cpi/cpifact4.htm

It comes as no surprise that non-fee schedule states tend to be among the jurisdictions with elevated rates of inflation. Most interestingly, for several fee schedule states, the average level of inflation over the studied time period of nearly 11 years is close to zero.

Chart 2 presents for the 36 jurisdictions the average rate of price inflation (All CPT), alongside the average rate of utilization increase. Clearly, for most jurisdictions, the rate of inflation exceeds the rate of utilization increase. At the same time, there appears to be no systematic relation between the rates of price level and utilization increases. The differences in the sum of price level and utilization increases between New Hampshire and Montana (on one side) and Texas and Oklahoma (on the other side) is remarkable, given that these averages cover a period of nearly 11 years.

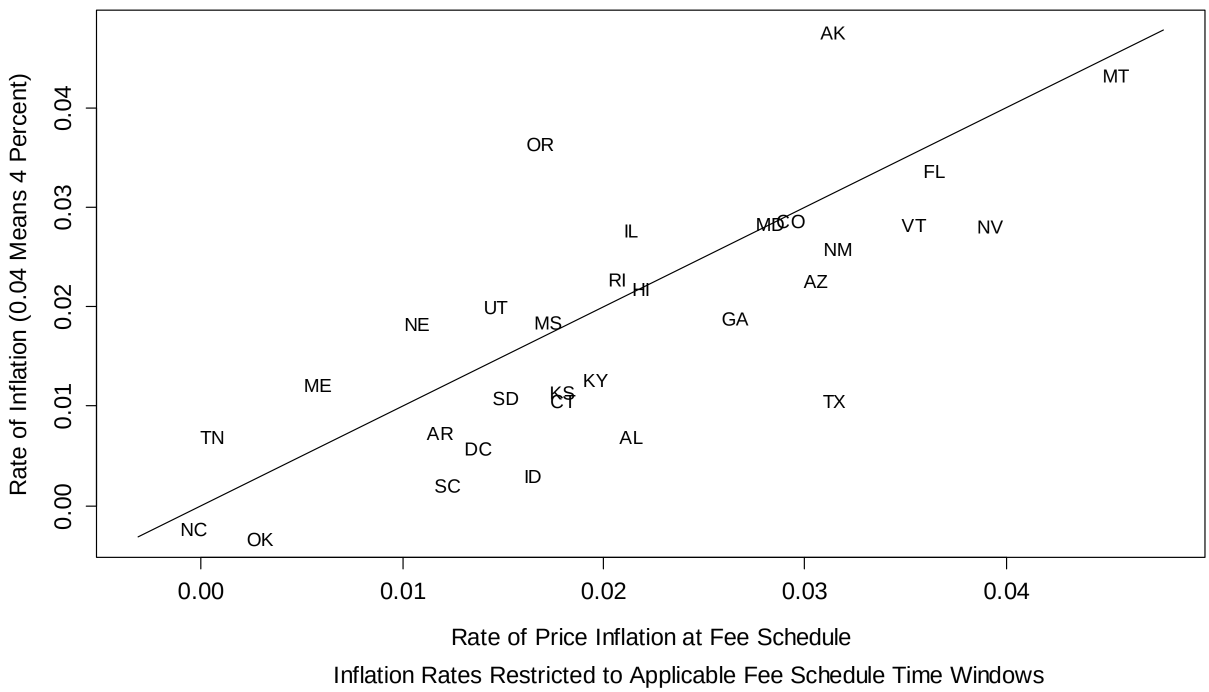
Chart 2: Utilization Increase vs. Price Inflation, February 2000–December 2010, CAGR



Finally, for the 31 fee schedule jurisdictions, Chart 3 plots the rate of price inflation at reimbursed amounts (All CPT) against the rate of inflation implied by the price ceilings laid down in the fee schedule (that is, the set of fixed-value MAR). Although there is a strong correlation between the two measures of price level increases, for some states, the two rates of inflation differ materially.

For Florida and Georgia, the average rate of price level increase at reimbursed amounts falls short of the average rate of increase at fee schedule, which results in a widening price departure, as documented below. Conversely, the price departure tends to shrink where the rate of price level increase at reimbursed amounts exceeds the rate of increase of the price level at MAR. Clearly, the price departure is bounded at zero (although reimbursement above fee schedule is possible in some jurisdictions). By implication, once the price level at reimbursed amounts reaches the price level at fee schedule, the rate of increase of the former can no longer exceed the rate of increase of the latter. Averaged equally weighted across the 31 states and over the analyzed 11 years, the price level at reimbursed amounts falls short of the price level at fee schedule by 10.7 percent.

Chart 3: Price Inflation vs. Fee Schedule, February 2000–December 2010, CAGR



4.2 Selected States

Florida and Georgia are chosen for a descriptive data exploration of the impact of fee schedule changes. The fee schedule trajectories of the two states are significantly different. Whereas in Florida, the price level implied by the MAR specified in the fee schedule has not risen (but rather, has declined) between a material fee schedule increase in 2005 and the end of the period analyzed (December 2010), in Georgia, which also recorded a significant fee schedule increase in 2005, the fee schedule increased each year. Notably, in both jurisdictions, the reimbursed amounts have increasingly fallen short of the MAR in the second half of the period analyzed.

4.2.1 Florida

Chart 4, top panel, depicts the Fisher price indexes of medical services at fee schedule and at reimbursed amounts (All CPT, which is the default price index).

For Florida, fee schedule information is not available to the authors before September 30, 2001. The first displayed price index at fee schedule pertains to December 2001, which is the third full month for which fee schedule information is available. For this first month, the price index at fee schedule and the two price indexes at reimbursed amounts are normalized to unity. If, subsequently, the price index at reimbursed amounts rises above the price index at fee schedule, this implies that since that first common month, on average, the inflation rate of the former exceeds the inflation rate of the latter (as measured by the compound annual growth rate). As mentioned, the price index at fee schedule may change from month to month, even as the fee schedule does not. This is because the distribution of transactions by place of service may vary over time.

Chart 4: Florida: Fee Schedule, Price Level, and Price Departure, January 2000–December 2010

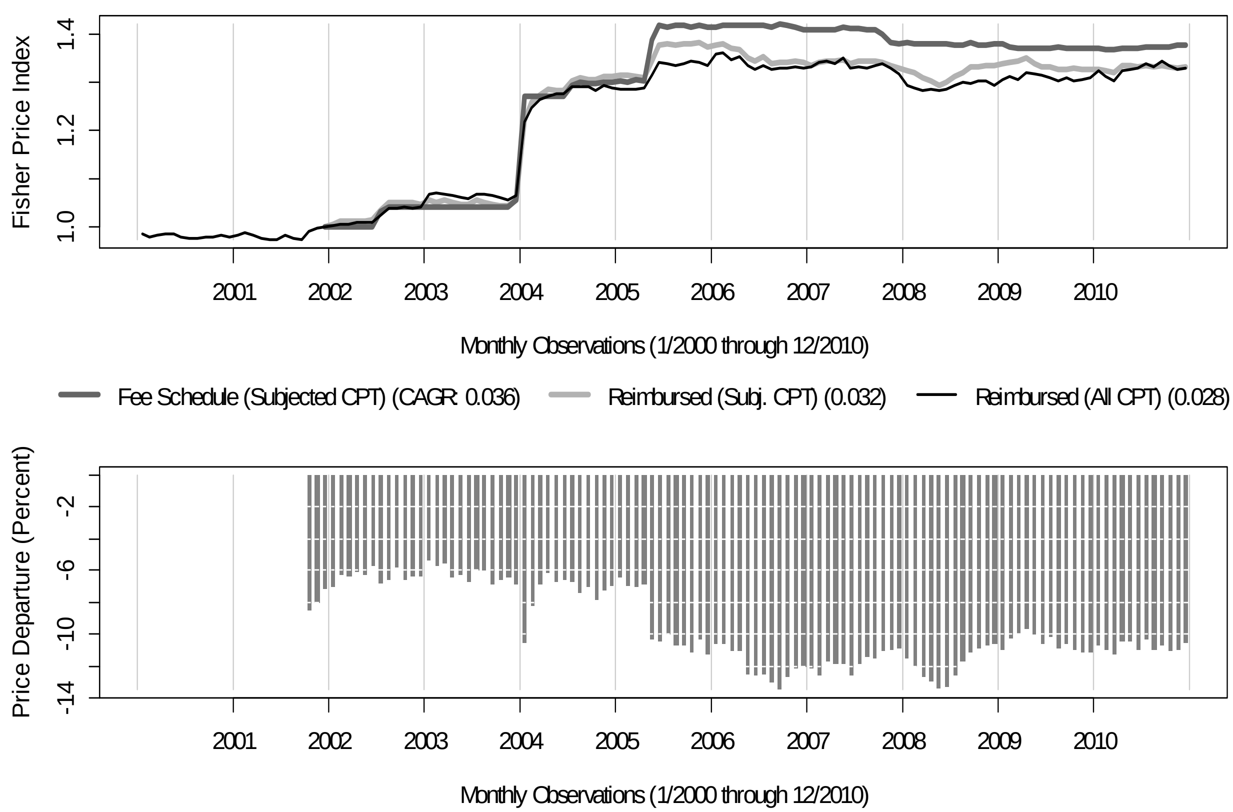


Chart 4, bottom panel, displays the price departure for All Categories, starting in October 2001, which is the first full month for which fee schedule information is available to the authors. As discussed, the price departure is calculated as the ratio of transaction volume at reimbursed amounts to transaction volume at fee schedule, minus 1. For a given fee schedule, an increase in the price index at reimbursed amounts tends to diminish the magnitude of the price departure, thereby shortening the depicted bars. Yet, as pointed out, changes in the price departure do not perfectly correlate with changes in the top panel of the vertical distances between the price index at fee schedule and the price indexes at reimbursed amounts.

Chart 4 reveals that from December 2001 through April 2005, the three price indexes (top panel) track each other closely. Then, following a fee schedule increase in May 2005, the two price indexes at reimbursed amounts rise less than what is implied by the fee schedule increase. As a result of the inelastic price level response to the fee schedule increase, the price departure (bottom panel) widens. The gap between the price index at fee schedule and the price indexes at reimbursed amounts remains in place through the end of the period analyzed (December 2011); consistent with the persistence of the inelastic price level response, the price departure remains elevated with no sign of mean reversion.

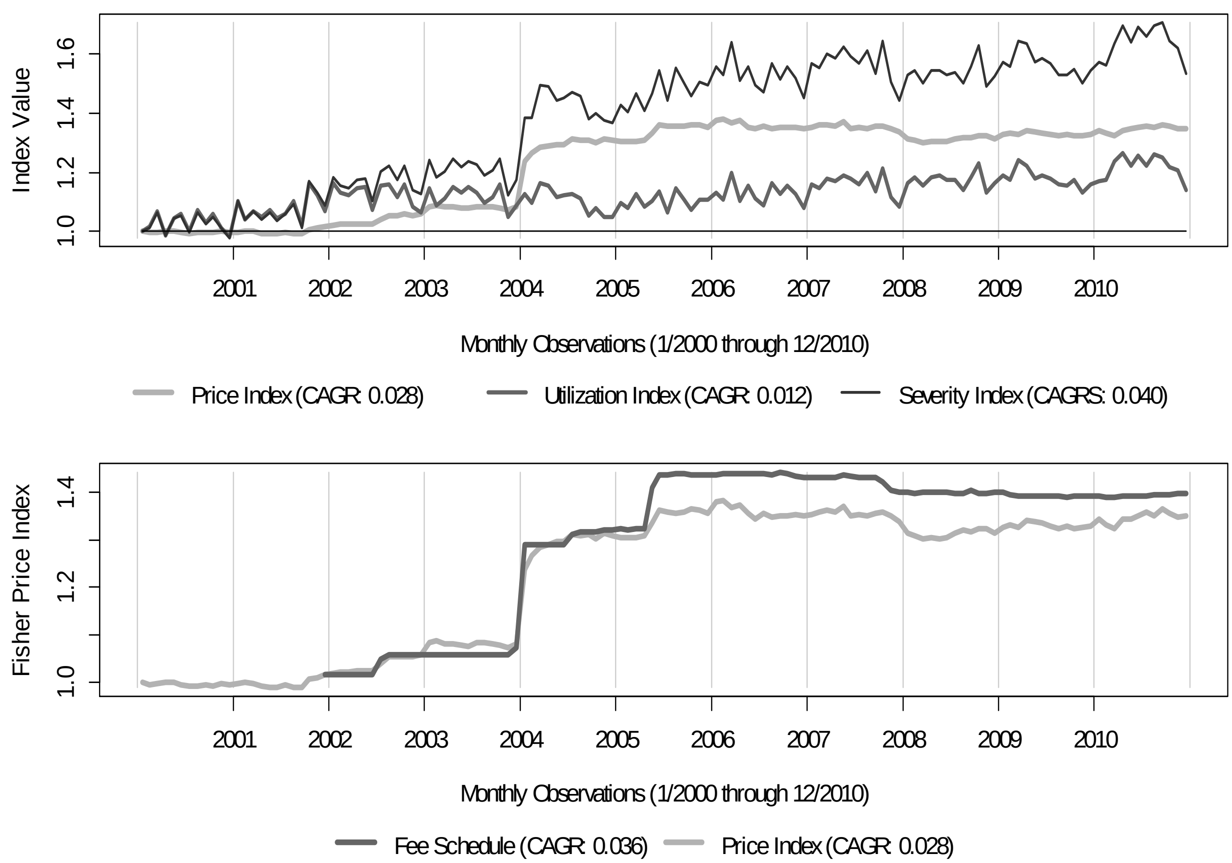
In October 2007, a fee schedule decrease in Florida manifests itself in a slight dip in the price index at fee schedule (Chart 4, top panel). This decrease in the price index at fee schedule is the first material change of this index since the fee schedule increase of May 2005. Similarly, in the three years following the fee schedule decrease, the price index at fee schedule remains nearly constant, thus implying a rate of inflation (at fee schedule) of approximately zero.

Over the time window of nearly 11 years (February 2001 through December 2011), the rate of inflation at fee schedule measures an annual 3.6 percent, on average. This compares to rates of inflation at reimbursed amounts of 2.8 percent (All CPT) and 3.2 percent (Subjected CPT).

Chart 5 depicts the utilization and severity indexes, alongside the price indexes at reimbursed amounts (all CPT codes) and at fee schedule. The indexes are normalized to unity in January 2000, except for the price index at fee schedule (bottom panel), which is normalized in its first month to the value of the price index at reimbursed amounts (which, in turn, is normalized to unity in January 2001). The utilization index is comparatively jagged, which translates into a similarly jagged severity index. The behavior of the utilization index around major fee schedule changes (as measured by major changes in the price index at fee schedule) reveals no perceptible link to the variation in MAR. The absence of a utilization effect in response to fee schedule changes will be confirmed in the statistical analysis below.

As shown in Chart 5, top panel, in Florida, the rate of inflation averages more than twice the rate of utilization growth over the period analyzed of nearly 11 years; whereas the rate of inflation measures 2.8 percent per annum, utilization records a 1.2 percent annual rate of increase.

Chart 5: Florida: Price Level, Utilization, and Severity, January 2000–December 2010



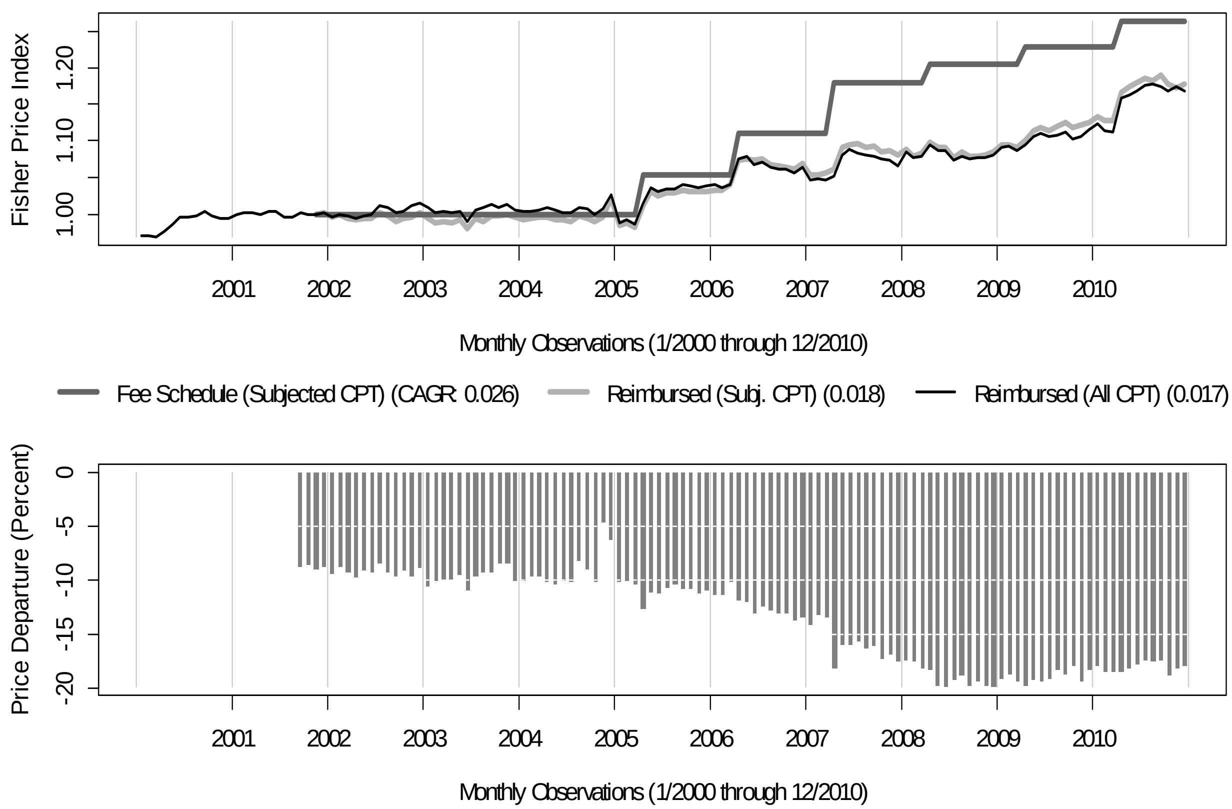
4.2.2 Georgia

Chart 6 displays for Georgia the price indexes at reimbursed amounts and at fee schedule, alongside the price departure. The first fee schedule information available to the authors bears an effective date of September 1, 2001. Consequently, the first price index values pertain to November 2001—for this month, all price indexes are normalized to unity. From November 2001 through March 2005, the price index at fee schedule holds steady, and the price indexes at reimbursed amounts (All CPT codes and Subjected CPT codes) remain largely flat.

Then, starting in 2005, there are annual fee schedule increases—the annual rate of inflation implied by the price index at fee schedule over the period 2006 through 2010 averages 4.0 percent.

By comparison, the annual rates of inflation of the price indexes at reimbursed amounts average only 2.4 percent (All CPT) and 2.7 percent (Subjected CPT). As a result of this extensive difference in the rates of inflation between the fee-schedule-imposed MAR and the reimbursed amounts, the price departure widens considerably. During the final three years of the analysis, the price level at reimbursed amounts is between 15 and 20 percent below the price level implied by the fee schedule price ceilings. Interestingly, the fee schedule kept increasing for many years in spite of an already significant price departure. As an informational item, for the final month of the analysis, the percentage (dollar) transaction volume that is subject to the fee schedule in Georgia equals 93.8 percent. (As mentioned, the service category Anesthesia is not included in the analysis.)

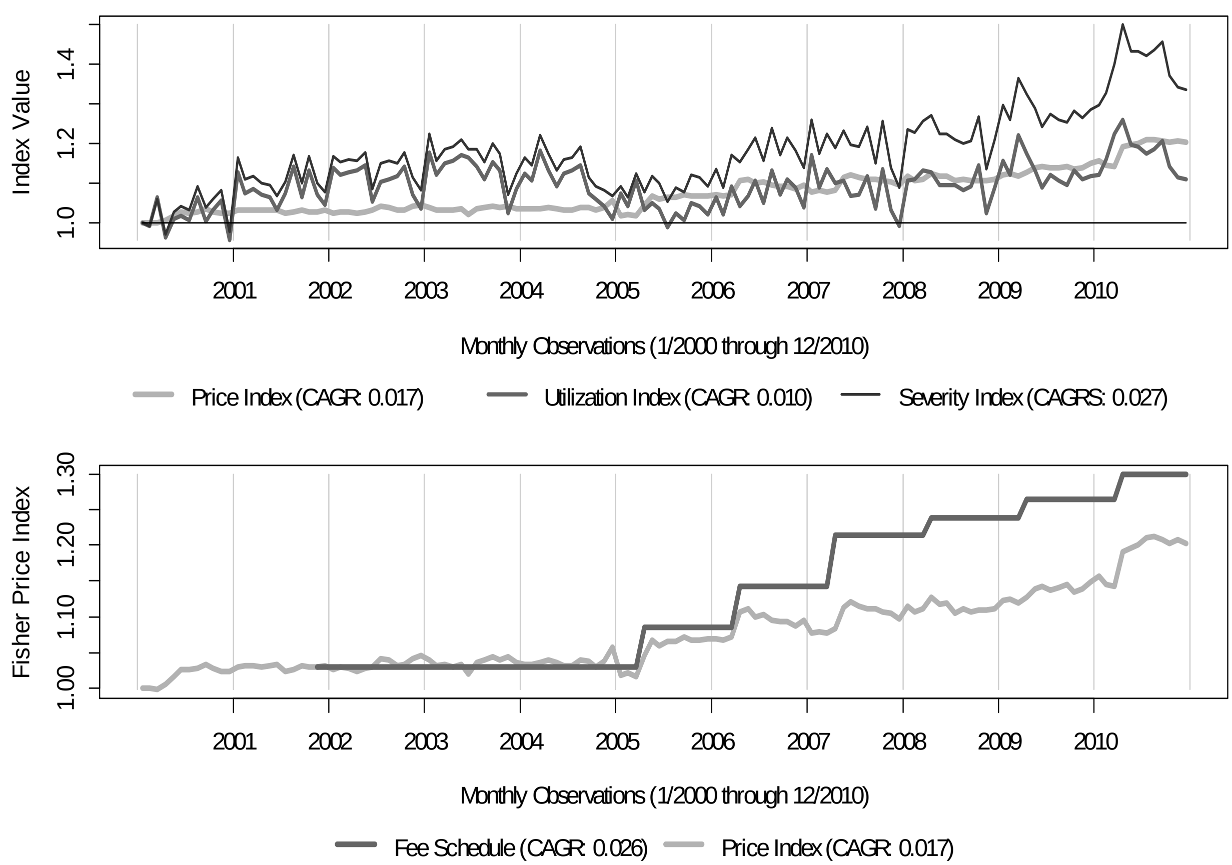
Chart 6: Georgia: Fee Schedule, Price Level, and Price Departure, January 2000–December 2010



Finally, Chart 7 presents for Georgia the utilization and severity indexes, alongside the price indexes at reimbursed amounts (All CPT) and at fee schedule. Here, too, the indexes are normalized to unity in January 2000, except for the price index at fee schedule (bottom panel), which is normalized in its first month to the value of the price index at reimbursed amounts (which, in turn, is normalized to unity in January 2001). Once again, the utilization index is comparatively jagged, thus translating into a jagged severity index. As with Florida, the behavior of the utilization index around major fee schedule changes (as measured by major changes in the price index at fee schedule) shows no clear link to the variation in the fee schedule.

As documented in Chart 7, in Georgia, over the nearly 11-year period analyzed, the average annual rate of inflation exceeds the rate of utilization growth by 70 percent (1.7 percent vs. 1.0 percent). The rate of utilization increase in Georgia is similar to Florida, where it equals 1.2 percent per annum.

Chart 7: Georgia: Fee Schedule, Price Level, and Price Departure, January 2000–December 2010



5. TIME SERIES MODELS

The impact of fee schedule changes is quantified in two time series models. Both models quantify the (logarithmic) rate of change in the severity index in response to fee schedule changes and an autonomous component (which represents drift). Prior to applying these models, stationarity of the rates of change in the price and utilization indexes is established. Then, the utilization index series of the individual jurisdictions are seasonally adjusted and, using these seasonally adjusted utilization index series, seasonally adjusted severity index series are calculated. All rates of growth (e.g., inflation rates) are expressed as differences in natural logarithms.

The impulse originating in the fee schedule change is quantified as the product of a Laspeyres index and the proportion of the transaction volume that is subject to fixed-value MAR. The Laspeyres index evaluates the quantities of the prior month at the prices of the current month (numerator) and the prices of the prior month (denominator)—this way, this price index isolates the price effect of a fee schedule change. If a fee schedule change occurs mid-month, then the impulse extends over two months. The transaction volume, which serves as a weight in the computation of the impulse, dates from the same month as the quantities of the corresponding Laspeyres index.

The response to an impulse may spread out over several months, as billing practices (and potentially, medical consumption) may take time to adjust. One way of modeling such lagged responses is to impose a specific functional form—this is to avoid the proliferation of regression coefficients, where impulses are highly correlated over time. On the other hand, imposing such a structure harbors the risk of predetermining a trajectory for the response that has little basis in the data. In this study, due to the sparseness of fee schedule changes, there is little potential for correlation among the covariates in an unstructured lag, which obviates the need for imposing a specific functional form.

The unstructured lag is given a length of 11 months. Allowing a full calendar year for the effect of a fee schedule change to manifest itself in the data ensures that seasonal effects (some of which may be present even after seasonally adjusting the data) do not adversely affect the estimated response. Further, fee schedule increases serve the realignment of prices with operating costs. If operating costs increase continually but the fee schedule is increased only once a year, then there may be temporary supply changes, which dissipate over the course of time.

A possible extension of the two models in future work is to allow for random effects in the parameters (across jurisdictions) by affording the model a multilevel structure.

The statistical models are Bayesian and estimated by means of MCMC (Markov-chain Monte Carlo simulation). The simulations are performed using JAGS (Just Another Gibbs Sampler, mcmc-jags.sourceforge.net), which is called from R (cran.r-project.org). JAGS is linked to R via proprietary software that makes use of parallelization. The JAGS code is displayed in Sections 7.8 and 7.9 of the appendix.

5.1 Stationarity

Inflation rates are known to have a high degree of persistence and may even follow a random walk. Similarly, the rate of utilization growth may be nonstationary. Before applying the statistical models, it is established that the rates of price inflation and utilization growth do not have unit roots.

The degree of integration is determined by means of the auto.arima procedure of the R package *forecast*, which was developed by Hyndman and Khandakar [7]. The auto.arima procedure allows for seasonality in the time series. Model selection was performed using the AICc information criterion, as suggested by Hurvich and Tsai [6]; no stepwise selection was permitted.

For every jurisdiction, the auto.arima procedure is applied to the time series of the rate of inflation at reimbursed amounts (All CPT) and the rate of utilization growth. Of the 31 inflation time series, only one is found to be nonstationary. By contrast, the utilization growth time series are all found to be integrated of order zero. Thus, in the following statistical models, the time series are uniformly treated as stationary.

5.2 Seasonal Adjustment

In many jurisdictions, the utilization index (and, hence, the severity index) exhibits a seasonal pattern. For this reason, the utilization index is seasonally adjusted using the X12-ARIMA software (version 0.3, 2011) of the U.S. Census Bureau. A seasonally adjusted severity index is then calculated as the product of the price index and seasonally adjusted utilization.

Interestingly, fee schedule changes exhibit a seasonal pattern as well, as demonstrated in Chart 8. Due to the seasonality of fee schedule changes, seasonally adjusting the utilization (and, by implication, the severity) index poses the risk of tempering the utilization responses to fee schedule changes. Then again, without seasonal adjustment, spurious correlation between seasonal variation in utilization and fee schedule changes may adversely affect the estimated utilization response to changes in MAR. In a sensitivity analysis, the main results of the analysis are compared to estimates from nonseasonally adjusted data.

5.3 Global Responses of Price, Utilization, and Severity

The first statistical model establishes the response to fee schedule changes of the rate of inflation (All CPT), the rate of utilization growth, and the rate of severity growth, disregarding state-level characteristics. The three growth rates are modeled simultaneously. A constraint ensures that the estimated rates of inflation and utilization growth add up to the estimated rate of severity growth. Further, the rates of inflation and utilization growth are given an intercept, which represents an autonomous growth component. This autonomous growth component represents drift in the price and utilization levels, which gives rise to drift in the severity level. As shown by Charts 4 and 6, the rate of price inflation at reimbursed amounts (All CPT) need not fully agree with the rate of price inflation at fee schedule, if only because not all CPT codes are subject to fixed-value MAR; besides, the price departure may widen or shrink over the analyzed 11 years. As to drift in utilization, Chart 2 demonstrates that the average rate of utilization growth is almost uniformly positive.

The three-equation model allows the responses to fee schedule changes to be asymmetric. As discussed, from a long-term perspective, the price levels at reimbursed amounts and at fee schedule must move in lockstep, which implies symmetry in the fee schedule responses (on the applied logarithmic scale). Then again, as discussed, the rate of inflation (at reimbursed amounts) may fall short of the rate of fee schedule increases for many years running.

Chart 8: Fee Schedule Changes by Calendar Month, February 2000–December 2010

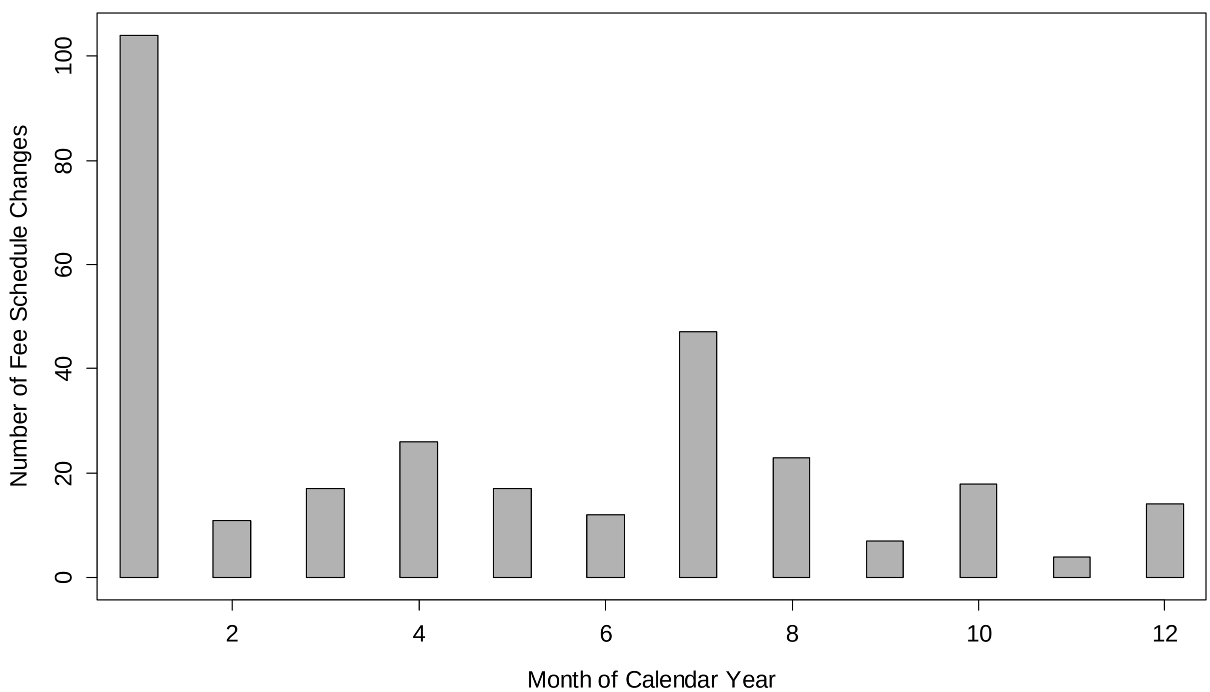


Chart 9 shows the estimated responses to fee schedule increases, displayed as a cumulative percentage change. In the chart, month 0 indicates the month of the fee schedule change. (Note that for mid-month fee schedule changes, due to prorating, the impulse extends to two months. Both prorated impulses start in month 0.) The chart shows that the price response is nearly completed in the third month following the fee schedule change. In month 11, the cumulative utilization response is essentially zero, thus leaving the price response as the sole contributor to the severity response. The cumulative severity response amounts to about 81 percent.

Chart 9: Price, Utilization, and Severity Responses to Fee Schedule Increases

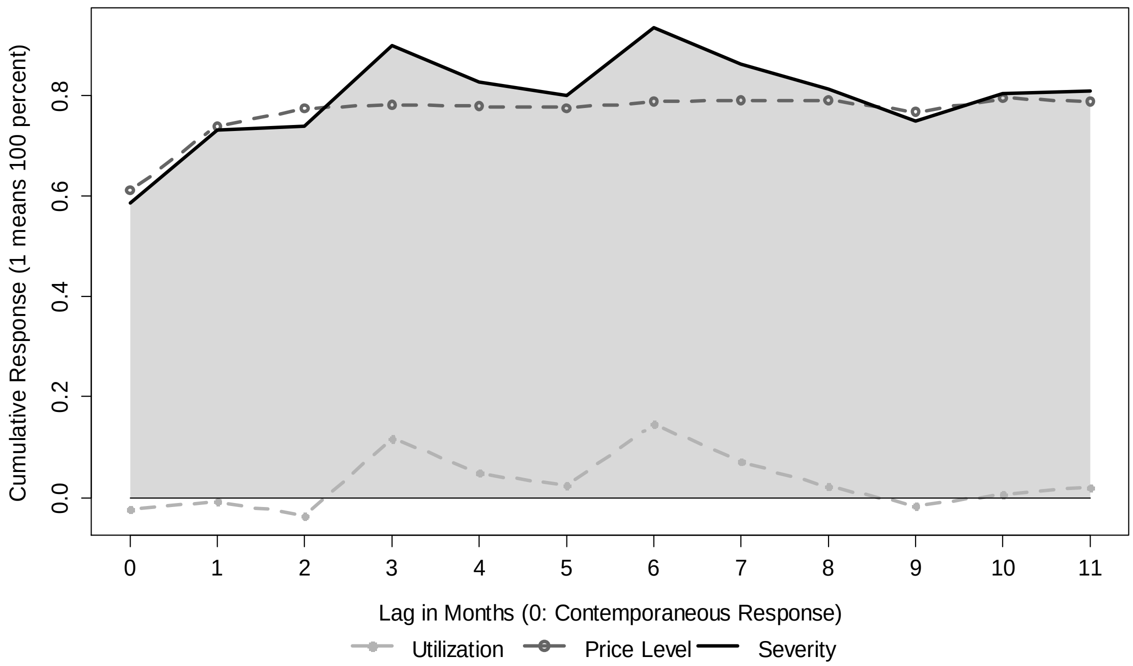


Chart 10 depicts the global responses to fee schedule decreases. Due to the comparatively small number of observations, the estimated cumulative responses are less smooth than for fee schedule increases. Here again, in month 11, the estimated severity response consists essentially of the price response only—the estimated cumulative severity response comes to about 45 percent. The chart reveals no lasting utilization increase (which would be indicated by a negative cumulative response). The width of the 80 percent credible intervals for the cumulative utilization response depicted in Chart 11 supports this conclusion, although a temporary effect cannot be ruled out.

Whereas the first statistical model offered a breakdown of the severity response into price level and utilization responses, the second statistical model attempts to deliver differentiated severity responses based on state-level characteristics. This model is applied to fee schedule increases only; this is because observations of fee schedule decreases are considered too sparse to render reliable estimates at such granularity.

Chart 10: Price, Utilization, and Severity Responses to Fee Schedule Decreases

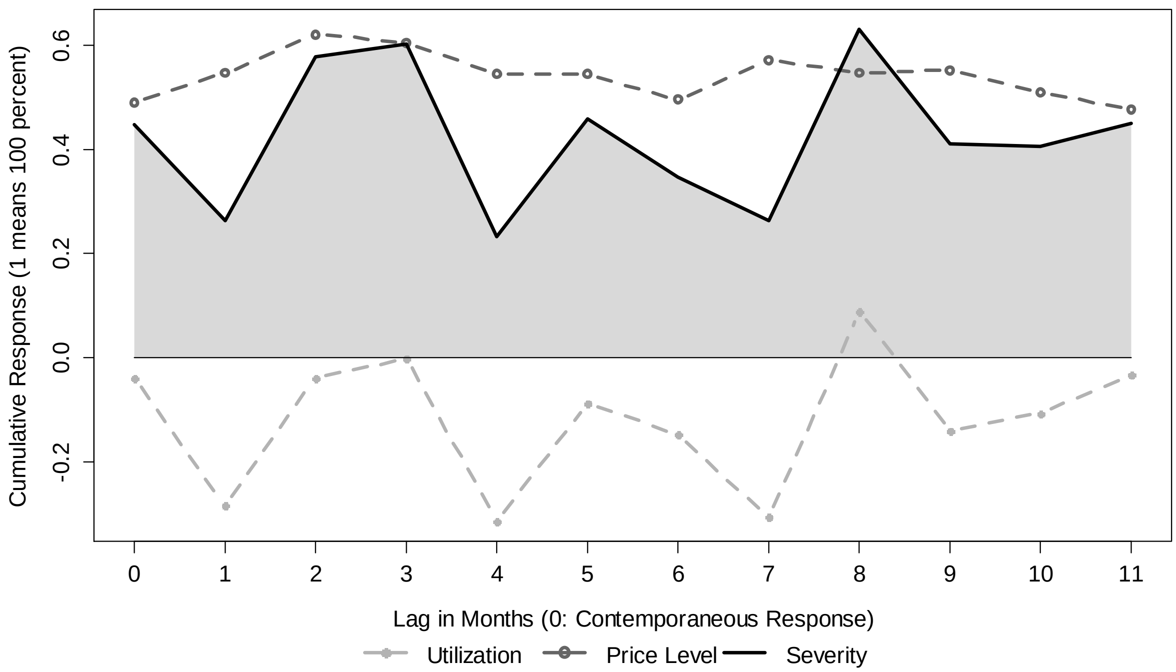
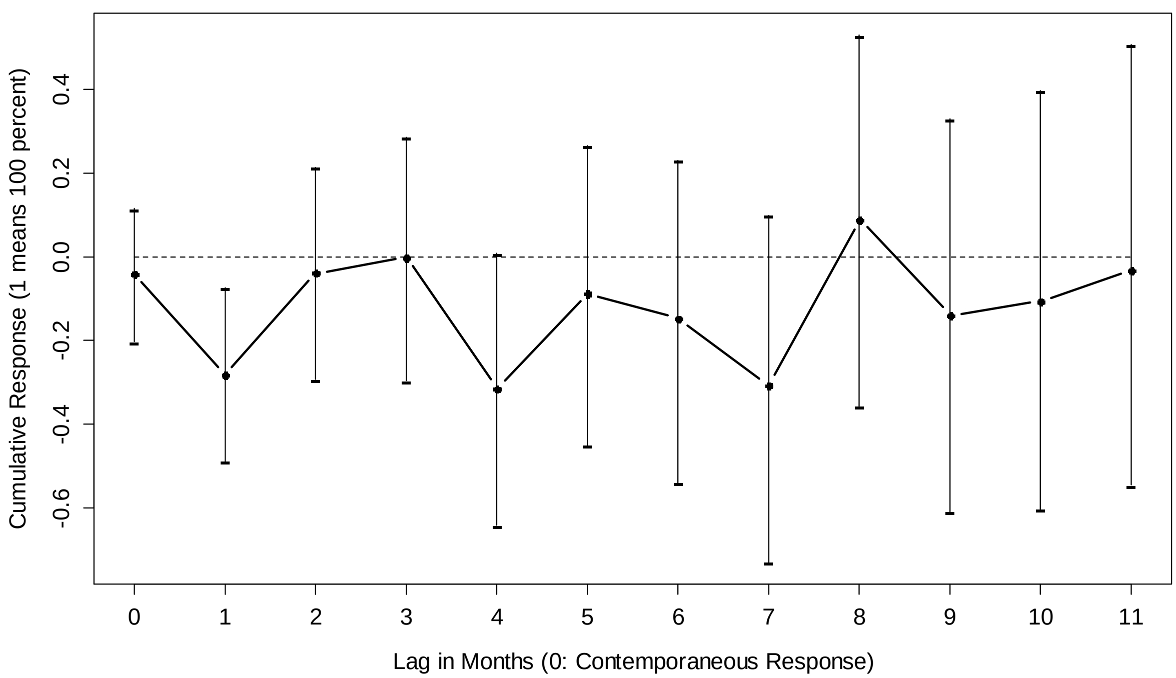


Chart 11: Utilization Response to Fee Schedule Decreases, 80 Percent Credible Intervals



Two alternative state-level characteristics are considered. First, the response to a fee schedule increase may depend on the magnitude of the price departure prior to the new MAR taking effect. Second, the magnitude of the response may be related to the price level (at reimbursed amounts) relative to neighboring jurisdictions. This price level relative to neighbors is calculated using a Lowe index based on the Star method; for details see Appendix 7.6.

An extensive price departure may be read as an indication that the prevailing fee schedule is comparatively high—here, as compared to the reimbursed amounts. As a result of a comparatively high fee schedule, further increases in MAR may elicit comparatively weak responses in the price level at reimbursed amounts. Similarly, a high price level at reimbursed amounts, as compared to neighboring jurisdictions, may render a comparatively small multiplier.

The statistical model is again a three-equation framework, where one equation represents the global response, a second equation captures the response differentiated by the price departure, and the third equation delivers multipliers that vary by the differential in the price level (at reimbursed amounts) relative to neighboring states. Specifically, the multipliers read:

$$\lambda \tag{4a}$$

$$\lambda \cdot \left(1+\kappa \cdot \left(\text{Price Departure}_i - \frac{1}{N}\sum_{i=1}^{N}\left(\text{Price Departure}_i\right)\right)\right) \tag{4b}$$

$$\lambda \cdot \left(1+\eta \cdot \left(\text{Price Level Relative to Neighbors}_i - \frac{1}{N}\sum_{i=1}^{N}\left(\text{Price Level Relative to Neighbors}_i\right)\right)\right) \tag{4c}$$

where the index refers to the fee schedule increase i in the total of N observations. When the price departure in the month prior to the fee schedule change equals the average price departure observed in those final months, then Equation (4b) reduces to Equation (4a) and, consequently, the multiplier equals the global response, λ. Similarly, when in the month preceding a fee schedule increase, the price level differential to neighboring states equals the average price level differential observed in the final months leading up to a fee schedule change, then the multiplier equals the global response.

Note that the price departure employed in the statistical analysis is defined differently than the price departure exhibited in Charts 4 and 6. In the charts, price departure is defined as the ratio of transaction volume at reimbursed amounts to transaction volume at fee schedule, minus 1. By contrast, in the statistical analysis, price departure is defined as the (natural) logarithm of the ratio of transaction volume at reimbursed amounts to transaction volume at fee schedule. Thus, adding 1 to the price departure concept employed in the charts and then taking the logarithm delivers the price departure concept used in the statistical model. Similarly, the price level relative to neighbors equals the (natural) logarithm of the mentioned Lowe price index.

Further, the price level differential relative to neighbors is not necessarily zero when averaged across all analyzed months (including those without fee schedule changes); this is because the Star method used in calculating the Lowe index affords more weight to larger states (that is, states with higher transaction volumes). As a result, a large state with a comparatively high price level may cause the majority of states to have a Lowe index less than unity (or, equivalently, a negative price differential).

The estimated coefficients of the multipliers read (after rounding to two decimal places):

$$0.76 \tag{5a}$$

$$\begin{aligned} &0.76 \cdot \left(1 + 1.25 \cdot \left(\text{Price Departure}_i + 0.10\right)\right) \\ = &0.76 \cdot \left(1.125 + 1.25 \cdot \text{Price Departure}_i\right) \end{aligned} \tag{5b}$$

$$\begin{aligned} &0.76 \cdot \left(1 - 1.17 \cdot \left(\text{Price Level Relative to Neighbors}_i + 0.05\right)\right) \\ \approx &0.76 \cdot \left(0.94 - 1.17 \cdot \text{Price Level Relative to Neighbors}_i\right) \end{aligned} \tag{5c}$$

The estimated regression coefficients support the hypothesized relation between the magnitude of the severity response to the fee schedule change and the extent of the price departure that prevailed in the month prior to the new MAR taking effect. The more pronounced the price departure is, the smaller the number is by which the global response of 0.76 is to be multiplied. Similarly, the hypothesized relation between the strength of the severity response to the fee schedule change and the price level (at reimbursed amounts) relative to neighboring jurisdictions is borne out by the regression results. The higher the price level relative to the neighbors, the weaker is the severity response to the fee schedule change. (Remember that the price level response accounts essentially for the entire severity response.)

Although the regression results concerning the influence of the price level relative to neighbors help improve the understanding of market behavior, the findings are of little value from an implementation standpoint. This is because the evaluation of a fee schedule change in any given state would require the computation of the Lowe price indexes for all 31 states used in this analysis. Further, the fee schedule of a jurisdiction not included in this analysis could not be evaluated.

As mentioned, if the policymaker reacts to market conditions, as measured by the price departure and the price level relative to neighbors, then a correlation arises between these covariates (on one side) and the error term in the respective regression equation (on the other side). As a result of this correlation, the influence of the covariates is underestimated (that is, the coefficients are biased toward zero). Although there are formal tests on the presence of such an endogeneity bias, in the context studied here, discerning whether policy makers react to market conditions is comparatively straightforward. If a narrow (extensive) price departure is causal to a fee schedule increase (decrease), then, on average, the price departure in the month prior to a fee schedule increase can be expected to be less extensive than the price departure in the month prior to a fee schedule decrease. Likewise, if a comparatively low (high) price level relative to neighboring jurisdictions contributes to a fee schedule increase (decrease), then the price level relative to neighbors can be expected to be lower prior to a fee schedule increase than in the time leading up to a fee schedule decrease.

On average, in the month prior to a fee schedule increase, the price departure (as defined in the statistical model) equals a negative 0.105. By comparison, the average price departure prior to a fee schedule decrease equals a negative 0.103. If fee schedule changes were driven by the price departure, then the price departure prior to fee schedule decreases would be more extensive, not less. From this perspective, the regression coefficient that reflects the influence of the price departure is not expected to be subject to an endogeneity bias.

Matters are different for the price level relative to neighbors. Prior to a fee schedule increase, on average, the log of the price level relative to neighbors equals a negative 0.048; this compares to a value of a negative 0.045 in the months prior to fee schedule decreases. Clearly, an endogeneity bias in the regression coefficient capturing the influence of the price level to neighbors cannot be ruled out.

When estimating the model using nonseasonally adjusted data, the following coefficients of the multiplier result (after rounding to two decimal places):

$$0.79 \tag{6a}$$

$$\begin{aligned} &0.79 \cdot \left(1 + 3.72 \cdot \left(\text{Price Departure}_i + 0.10\right)\right) \\ &= 0.79 \cdot \left(1.372 + 3.72 \cdot \text{Price Departure}_i\right) \end{aligned} \tag{6b}$$

$$\begin{aligned} &0.79 \cdot \left(1 - 0.89 \cdot \left(\text{Price Level Relative to Neighbors}_i + 0.05\right)\right) \\ &\approx 0.79 \cdot \left(0.96 - 0.89 \cdot \text{Price Level Relative to Neighbors}_i\right) \end{aligned} \tag{6c}$$

Based on nonseasonally adjusted data, the magnitude of the global response exceeds the one obtained for seasonally adjusted data. It appears that some of the seasonal variation in the data may be related to the seasonal variation of fee schedule changes. Interestingly, the leverage afforded to the price departure is now considerably higher, whereas the influence of the price level relative to neighbors is somewhat diminished.

Clearly, the results that were obtained from seasonally adjusted data are to be preferred. This is because the seasonal adjustment applies to the utilization index only (and not to the price index). Seasonal variation in utilization manifests itself in seasonal variation of the severity index and, thus, may distort the estimated influence of fee schedule changes in a meaningful way.

When it comes to the evaluation of fee schedule changes based on Equation (5b), two limitations apply. These two limitations arise from the constraint that the price departure cannot turn positive, notwithstanding the possibility of reimbursement above MAR for isolated CPT codes within the legal limits. First, in the event of a fee schedule increase, the response in the price level shall not exceed the impulse by an amount so large as to engender a positive price departure. A sufficient condition for this constraint to hold is that the multiplier displayed in Equation (5b) does not exceed unity. Inserting the least extensive price departure—a departure of zero—into Equation (5b) delivers a multiplier of 0.855, which satisfies the constraint. By implication, the highest possible multiplier in the context of a fee schedule increase equals 85.5 percent.

Second, in the event of a fee schedule decrease, the price level must follow the fee schedule down at least to the extent necessary for maintaining a nonpositive price departure. The only observation in the data set where the fee schedule decrease is close to having the potential of completely eliminating the existing price departure is Tennessee. In March 2008, the fee schedule reduction amounts to a 17.0 percent price level decrease (Subjected CPT). (This is prior to weighting the fee schedule change with the volume of subjected CPT codes to arrive at the impulse.) In the month prior to the fee schedule decrease, the price departure equals a negative 26.2 percent, based on the definition employed in the charts. (The fee schedule takes effect on March 4, 2008.) More than half of the original price departure withstands the fee schedule decrease.

Caution must be applied when extrapolating regression results beyond the observed data cloud. Some fee schedule increases occur in the context of a very small price departure. Thus, a price departure close to zero is within the realm of the observed. The most extensive price departure observed in the month prior to a fee schedule increase equals a negative 0.31 (as defined in the statistical model) or a negative 27 percent (as defined in the charts).

6. CONCLUSION

This study of the impact of physician fee schedule in the context of workers compensation is the first of its kind. In executing this study, new methodology was developed. Most important, a monthly utilization index was devised, which serves as a measure of utilization of medical services provided by physicians to workers compensation claimants.

Several findings may question conventional wisdom. First, the rate of price inflation of medical services provided by physicians in the context of workers compensation tends to fall short of the rate of price inflation of the Medical Care component of the CPI. Second, fee schedule changes have no lasting utilization effect.

Among the anticipated findings are the higher-than-average rate of inflation in jurisdictions without physician fee schedules and the continued increase in utilization in nearly all jurisdictions.

A simple formula for the evaluation of changes to physician fee schedules was devised. By using the relative difference in the price levels at fee schedule and at reimbursed amounts as a measure of market conditions, the formula is capable of factoring in state-specific information; further, this formula does not require information from jurisdictions other than the one that is evaluated.

The response to fee schedule changes is characterized by the absence of a utilization effect and by an inelastic price level response. To the degree that fee schedule changes follow changes in operating costs, no change in medical practice should be expected. In the long run, the price level at reimbursed amounts can be anticipated to move in lockstep with the fee schedule, at least with respect to the set of CPT codes that are subject to fixed-value MAR. Then again, the data show that the price levels at fee schedule and at reimbursed amounts may diverge for sustained periods of time.

7. APPENDIX

7.1 List of States Analyzed

For 36 jurisdictions, rates of change were computed for the level of reimbursed amounts (resulting in the rate of inflation), the level of utilization, and the level of severity. This list of states reads AK, AL, AR, AZ, CO, CT, DC, FL, GA, HI, IA, ID, IL, IN, KS, KY, MD, ME, MO, MS, MT, NC, NE, NH, NM, NV, OK, OR, RI, SC, SD, TN, TX, UT, VA, and VT. Further, the Lowe index of the price level at reimbursed amounts was computed for the entire set of 36 jurisdictions.

Of these 36 jurisdictions, 5 do not have fee schedules in place during the period analyzed (2000–2010): IA, IN, MO, NH, and VA. As such, 31 states are used in the statistical model.

Of the 31 jurisdictions for which fee schedule information is available, 3 introduced a fee schedule during the period analyzed: TN (fee schedule became effective on July 1, 2005), IL (February 1, 2006), and ID (April 1, 2006). The following set of 14 jurisdictions had fee schedules in place for the entire period analyzed, but fee schedule information was not available to the researchers prior to the fee schedule effective dates listed in parentheses: NC (March 1, 2000), AL (March 15, 2000), OR (April 1, 2000), CT (May 1, 2000), NV (May 1, 2000), AR (May 15, 2000), NE (June 15, 2000), RI (July 1, 2000), SD (July 19, 2000), UT (January 1, 2001), VT (January 1, 2001), GA (September 1, 2001), FL (September 30, 2001), and TX (September 1, 2002).

Following a fee schedule introduction or the first availability of an existing fee schedule, the first month that enters the statistical analysis is the third complete month following the fee schedule introduction or fee schedule availability, respectively. The first growth rate analyzed is the rate of change that pertains to the month following this third complete month. Note that in the statistical analysis, the price departure is applied with a time lag of one month, which implies that the first price departure entering the statistical analysis belongs to the mentioned third complete month.

In charts, the first value of the price index at fee schedule refers to the third complete month; this also applies to the price index at reimbursed amounts that comprises the CPT codes that are subject to the fee schedule. The first displayed price departure pertains to the first complete month following a fee schedule introduction or fee schedule availability.

7.2 Recoding Units of Service

For some medical services, the supplied quantity of service is measured in number of minutes. Yet, the units of service, as laid down in the fee schedule, refer to the number of time intervals, where a time interval is defined as a multiple of minutes. Medical services where the units of service are defined in such a way are common in Anesthesia (not covered in this study) and Physical Medicine.

In the data set employed, there were transaction records that appeared to report the number of minutes instead of the units of service. Before subjecting these transactions to the data-cleansing tools discussed below, an algorithm was applied for the purpose of recoding the reported units of service, if necessary. First, if the reported units equaled 9 or less, the reported value was left in place. Second, if the number of units reported was a multiple of 15, it was assumed that this reported value refers to the number of minutes. Third, if the number of units reported was not a multiple of 15 but was a multiple of 10, it was again assumed that the reported value reflects the number of minutes. Finally, if the reported value was greater than 9 but was not a multiple of 15 or 10, the units of service were treated as unknown—this way, the transaction is excluded from the computation of the price fences detailed below, yet is submitted to the associated outlier detection and management.

7.3 Outlier Detection

Outlier detection is undertaken independently for each jurisdiction and calendar year. At the center of the outlier detection approach is the schematic plot developed by Tukey [11]. The algorithm associated with this plot applies to reimbursement per unit of service on the natural log scale. Percentiles are indicated by the letter p; upper case format indicates a percentile on the raw scale, whereas the lower case points to the logarithmic scale. Outlier detection is performed on the level of the medical service and on the level of the service category. Price fences are defined for the purpose of restating prices and quantities.

For a given medical service, if $P_{75} \neq P_{25}$, then the price fences are set to $p_{75} + 0.6$ and $p_{25} - 0.6$. Exceptions are transactions for which the paid-to-submit ratio is greater than 0.5; in these cases, the lower fence is set to $p_{25} - 0.7$. Conversely, if $P_{75} = P_{25}$, then the price fences are set to $p_{85} + 0.2$ and $p_{15} - 0.2$. Transactions that fall inside the price fences of a given medical service make it into the calculation of the price fences at the level of the service category. The upper fence at the level of the service category is defined as the maximum of $p_{90} + 0.5$ (where the percentile is based on the service category) and the maximum of the upper fences across all medical services (in that service category) that register at least 20 transactions in the applicable calendar year. Similarly, the lower fence at the level of the service category is defined as the minimum of $p_{10} - 0.5$ and the minimum of the lower fences across all medical services that account for at least 20 transactions.

Transactions that are located inside the fences, both at the level of the medical service and the level of the service category, remain unedited. These transactions enter the computation of the average price per unit of service and the median units of service per transaction for the applicable medical service in a given month.

Medical services that register less than 12 records (in a given year and jurisdiction) are excluded from the analysis.

7.4 Outlier Management

Outlier management is undertaken on a calendar year basis, in keeping with the price fences, which were computed from data for the calendar year.

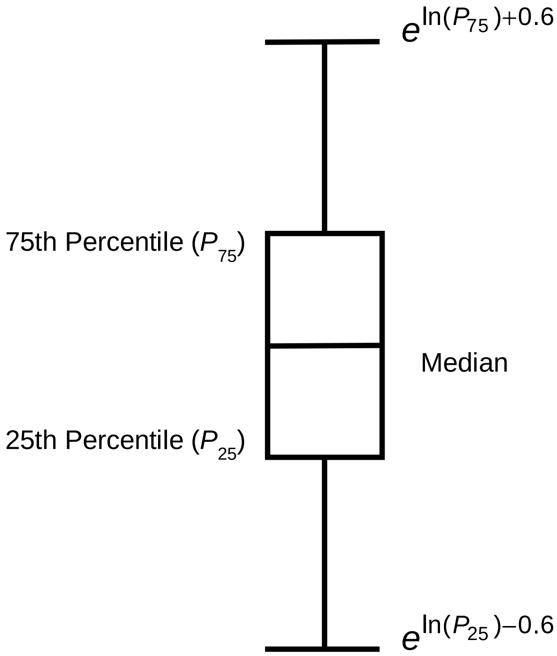
Transactions that come with prices per unit of service above the upper fences of the applicable medical service or the overarching service category are reset to the mean price of this medical service; the units of service associated with these transactions remain unaltered.

Transactions with prices per unit of service below the lower fences of the respective medical service or the applicable service category have their units of service restated based on the following process (which also applies to transactions with unknown units). First, the number of units is set to the median number of units per transaction for this medical service; this median number of units is typically unity. Then, the price is recalculated based on the units of service thus restated. If this recalculated price falls below the lower fences of the medical service or the service category, then the units of service associated with this transaction are set to unity. Then again, the price is recalculated. If this price still falls below any of the two applicable lower price fences, then the record is discarded as a nuisance transaction. Conversely, if, during this iterative process, the recalculated price exceeds any of the two upper fences, then this price is reset to the mean of the applicable medical service and no further restatement is done.

7.5 Tukey's Schematic Plot (Boxplot)

Chart A.1 depicts Tukey's schematic plot, also known as box-and-whiskers plot or box plot. The objective of the box plot is to report major location parameters (such as the median and the 25th and 75th percentiles) and to identify outliers. In this graph, the hinges that define the upper and lower limits of the box identify the IQR (inter-quartile range), which comprises 50 percent of the data. For the purpose of this study, the fences are defined on the logarithmic scale. The upper fence signifies the sum of the 75th percentile and 0.6 on the natural log scale (which corresponds to multiplying the 75th percentile on the raw scale by approximately 1.8). The lower fence equals the difference between the 25th percentile and 0.6 on the log scale (which amounts to dividing the 25th percentile on the raw scale by approximately 1.8). Values beyond the fences are considered outliers.

Chart 9: Tukey's Schematic Plot (Box Plot)



7.6 Multilateral Purchasing Power Parities

Price level differences across jurisdictions were calculated based on the concept of multilateral purchasing power parities. Technically, this index is a Lowe index, as discussed in Hill [5]. When calculating this index for the covariate *Price Level Relative to Neighbors*, the neighbors of a given state were defined based on Census divisions and, where applicable, regions (for details, see www.census.gov/econ/census07/www/geography/regions_and_divisions.html). The categories for the 36 states included in the Lowe index are:

New England: CT, ME, NH, RI, VT
North Central: IA, IL, IN, KS, MO, NE, SD
South Atlantic: DC, FL, GA, MD, NC, SC, VA
South Central: AL, AR, KY, MS, OK, TN, TX
Mountain: AZ, CO, ID, MT, NM, NV, UT
Pacific: AK, HI, OR

As calculated in the context of multilateral cross-jurisdiction price level comparisons, the Lowe index is constructed on a monthly basis, using the Star method. The Star is defined as comprising the transactions of all 36 jurisdictions. For any given jurisdiction, the equivalent to the international concept of the purchasing power parity is calculated as the ratio of (1) the quantities of this jurisdiction evaluated at its own prices to (2) these quantities evaluated at the prices of the Star. Note that in any given month, the denominator of this ratio may differ across jurisdictions as every jurisdiction may have its unique set of quantities of medical services. Hence, when taking the ratio of the Lowe indexes of two jurisdictions, the two denominators generally do not cancel.

The Lowe index calculated using the Star method allows for bilateral price level comparisons between the jurisdiction in question and the Star. More importantly, the ratio of the Lowe indexes of two jurisdictions provides a reading of their relative price levels. Similarly, the ratio of the Lowe index of a jurisdiction to the average Lowe index of the region provides a gauge of the price level relative to the neighbors (which are defined as belonging to the same geographical region, as detailed above). For the purpose of comparing the price level of a jurisdiction to its neighbors, the price level of the applicable region was calculated as an equally weighted average (which includes the price level of the pertinent jurisdiction). The concept of the equally weighted average treats the price level of every jurisdiction as a draw from the same underlying price level distribution, regardless of the size of the jurisdiction.

7.7 Hypothetical Numerical Example of a Fee Schedule Increase

The effect of a fee schedule change is quantified by the product of the impulse and the response to this impulse, called the multiplier. The multiplier used in this example rests on the original regression coefficients, rounded to one decimal place. (In Equation (4b), the original regression coefficients are rounded to two decimal places.)

First, a Laspeyres index is calculated. The numerator of this index consists of a weighted average of the new MAR, where the weights are the quantities of medical services consumed in the month prior to the fee schedule increase. The denominator of this index consists of a weighted average of the old MAR; the weights are the same as in the numerator. Note that only CPT codes with fixed-value MAR enter this computation. In a second step, for the month prior to the fee schedule change, the proportion of the dollar volume of medical services subject to the (then effective) fee schedule is calculated. Here, too, only CPT codes with a fixed value MAR are considered. The Laspeyres index and the proportion of volume, taken together, deliver the impulse. In a third step, the multiplier is applied to the impulse. As a numerical example, let us assume that the Laspeyres index equals 1.1 and that 90 percent of the volume is subject to MAR; for a multiplier of 0.8, the response equals $exp(log(1.1) \times 0.9 \times 0.8) - 1 \approx (1.1 - 1) \times 0.9 \times 0.8 \approx 0.072$ or, equivalently, 7.2 percent.

The multiplier of 0.8 represents a global concept, which applies to the analyzed jurisdictions on average. The multiplier can be refined using state-level data, thus accommodating state-specific conditions. This refinement can be accomplished in two (mutually exclusive) ways. First, the price departure can be factored into the multiplier. Assuming a price departure (as defined in the charts) of minus 5 percent (which implies that in this jurisdiction, based on a volume-weighted average, the reimbursed amount is 5 percent below MAR), the multiplier reads $0.8 \times (1.1 + 1.2 \times \log(1 + (-0.05))) \approx 0.83$.

Second, the price level in the state relative to its neighbors can be considered. Assuming that the Lowe index of the jurisdiction at hand equals 0.9 (which implies that in this state, compared to the neighboring jurisdictions, the price level of medical services provided by physicians is 10 percent lower), then the multiplier equals $0.8 \times (0.9 - 1.2 \times \log(0.9)) \approx 0.77$.

7.8 JAGS Code: First Three-Equation Model

```
model

{

for (j in 1:K){ #states
    for (i in t[j]:T){ #time
        delta.price[i,j] ~ dnorm(mu.price[i,j],tau.price[j])
        delta.quant[i,j] ~ dnorm(mu.quant[i,j],tau.quant[j])
        delta.sever[i,j] ~ dnorm(mu.sever[i,j],tau.sever[j])

        Y[i,j] ~ dsum(delta.pred.price[i,j],delta.pred.quant[i,j],delta.pred.sever[i,j])
        }
    }

for (j in 1:K){ #states
    for (i in 1:T){ #time

        delta.pred.price[i,j] ~ dnorm(mu.price[i,j],tau.price[j])
            mu.price[i,j] <- alpha.price[j] + sum(response.price[i,j,])

        delta.pred.quant[i,j] ~ dnorm(mu.quant[i,j],tau.quant[j])
            mu.quant[i,j] <- alpha.quant[j] + sum(response.quant[i,j,])

        delta.pred.sever[i,j] ~ dnorm(mu.sever[i,j],tau.sever[j])
            mu.sever[i,j] <- -(mu.price[i,j] + mu.quant[i,j])

        response.price[i,j,1] <- response.cat.price[i,j,L+1,1]
        response.price[i,j,2] <- response.cat.price[i,j,L+1,2]
        response.quant[i,j,1] <- response.cat.quant[i,j,L+1,1]
        response.quant[i,j,2] <- response.cat.quant[i,j,L+1,2]

        for (k in 1:(L+1)){

            response.cat.price[i,j,k,1] <- inprod(beta.price[1:k],pricemaronlykappa.price[i,j,1:k,1])
            response.cat.price[i,j,k,2] <- inprod(zeta.price[1:k],pricemaronlykappa.price[i,j,1:k,2])
            response.cat.quant[i,j,k,1] <- inprod(beta.quant[1:k],pricemaronlykappa.quant[i,j,1:k,1])
            response.cat.quant[i,j,k,2] <- inprod(zeta.quant[1:k],pricemaronlykappa.quant[i,j,1:k,2])

            #price impulse for price response
            pricemaronlykappa.price[i,j,k,1] <- pricemaronlydelta[i,j,k,1]
            pricemaronlykappa.price[i,j,k,2] <- pricemaronlydelta[i,j,k,2]

            #price impulse for quantity response
            pricemaronlykappa.quant[i,j,k,1] <- pricemaronlydelta[i,j,k,1]
            pricemaronlykappa.quant[i,j,k,2] <- pricemaronlydelta[i,j,k,2]

            }
        }
    }

for (i in 1:K){ #number of states
    alpha.price[i] ~ dnorm(0,1.0E-3)
    alpha.quant[i] ~ dnorm(0,1.0E-3)

    tau.price[i] <- pow(sigma.price[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)
    tau.quant[i] <- pow(sigma.quant[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)
    tau.sever[i] <- pow(sigma.sever[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)

    sigma.price[i] ~ dunif(0,4) #<- sqrt(2)/tau.price[i] #double exponential errors
    sigma.quant[i] ~ dunif(0,4) #<- sqrt(2)/tau.quant[i] #double exponential errors
    sigma.sever[i] ~ dunif(0,4) #<- sqrt(2)/tau.sever[i] #double exponential errors
    }
```

```
for (i in 1:(L+1)){ #L: number of lags
    beta.price[i] ~ dnorm(0,1.0E-3)
    beta.quant[i] ~ dnorm(0,1.0E-3)

    zeta.price[i] ~ dnorm(0,1.0E-3)
    zeta.quant[i] ~ dnorm(0,1.0E-3)

    beta.cum.price[i] <-  sum(beta.price[1:i])
    beta.cum.quant[i] <-  sum(beta.quant[1:i])
    beta.cum.sever[i] <-  sum(beta.price[1:i])+sum(beta.quant[1:i])

    zeta.cum.price[i] <-  sum(zeta.price[1:i])
    zeta.cum.quant[i] <-  sum(zeta.quant[1:i])
    zeta.cum.sever[i] <-  sum(zeta.price[1:i])+sum(zeta.quant[1:i])
    }

}
```

7.9 JAGS Code: Second Three-Equation Model

```
model

{

for (j in 1:K){ #states
    for (i in t[j]:T){ #time
        delta.sever.x[i,j] ~ dnorm(mu.sever.x[i,j],tau.sever.x[j])
        delta.sever.y[i,j] ~ dnorm(mu.sever.y[i,j],tau.sever.y[j])
        delta.sever.z[i,j] ~ dnorm(mu.sever.z[i,j],tau.sever.z[j])
        }
    }

for (j in 1:K){ #states
    for (i in 1:T){ #time

        delta.pred.sever.x[i,j] ~ dnorm(mu.sever.x[i,j],tau.sever.x[j])
            mu.sever.x[i,j] <- alpha.sever[j] + sum(response.sever.x[i,j,])

        delta.pred.sever.y[i,j] ~ dnorm(mu.sever.y[i,j],tau.sever.y[j])
            mu.sever.y[i,j] <- alpha.sever[j] + sum(response.sever.y[i,j,])

        delta.pred.sever.z[i,j] ~ dnorm(mu.sever.z[i,j],tau.sever.z[j])
            mu.sever.z[i,j] <- alpha.sever[j] + sum(response.sever.z[i,j,])

        response.sever.x[i,j,1] <- response.cat.sever.x[i,j,L+1,1]#P.sever,1]
        response.sever.x[i,j,2] <- response.cat.sever.x[i,j,L+1,2]#Q.sever,2]

        response.sever.y[i,j,1] <- response.cat.sever.y[i,j,L+1,1]#P.sever,1]
        response.sever.y[i,j,2] <- response.cat.sever.y[i,j,L+1,2]#Q.sever,2]

        response.sever.z[i,j,1] <- response.cat.sever.z[i,j,L+1,1]#P.sever,1]
        response.sever.z[i,j,2] <- response.cat.sever.z[i,j,L+1,2]#Q.sever,2]

        for (k in 1:(L+1)){

            response.cat.sever.x[i,j,k,1] <- inprod(beta.sever[1:k],pricemaronlydelta[i,j,1:k,1])
            response.cat.sever.x[i,j,k,2] <- inprod(zeta.sever[1:k],pricemaronlydelta[i,j,1:k,2])

            response.cat.sever.y[i,j,k,1] <- inprod(beta.sever[1:k],pricemaronlykappa.sever.y[i,j,1:k,1])
            response.cat.sever.y[i,j,k,2] <- inprod(zeta.sever[1:k],pricemaronlykappa.sever.y[i,j,1:k,2])

            response.cat.sever.z[i,j,k,1] <- inprod(beta.sever[1:k],pricemaronlykappa.sever.z[i,j,1:k,1])
            response.cat.sever.z[i,j,k,2] <- inprod(zeta.sever[1:k],pricemaronlykappa.sever.z[i,j,1:k,2])

            #price impulse for severity response (departure)
            pricemaronlykappa.sever.y[i,j,k,1] <- pricemaronlydelta[i,j,k,1] * (1 + chi.sever[1] *
                                                  (log(departure[i,j,k]) - departure.log.mean[1])
            pricemaronlykappa.sever.y[i,j,k,2] <- pricemaronlydelta[i,j,k,2] #* (1 + psi.sever[1] *
                                                  (log(departure[i,j,k]) - departure.log.mean[2])

            #price impulse for severity response (starindex)
            pricemaronlykappa.sever.z[i,j,k,1] <- pricemaronlydelta[i,j,k,1] * (1 + nu.sever[1] *
                                                  (log(starindex[i,j,k]) - starindex.log.mean[1])
            pricemaronlykappa.sever.z[i,j,k,2] <- pricemaronlydelta[i,j,k,2] #* (1 + xi.sever[1] *
                                                  (log(starindex[i,j,k]) - starindex.log.mean[2])


            }
        }
    }
```

```
chi.sever[2] <- 1 - chi.sever[1] * departure.mean[1]
psi.sever[2] <- 1 - psi.sever[1] * departure.mean[2]
nu.sever[2] <- 1 - nu.sever[1] * starindex.mean[1]
xi.sever[2] <- 1 - xi.sever[1] * starindex.mean[2]

for (i in 1:K){ #number of states
    alpha.sever[i] ~ dnorm(0,1.0E-3)

    tau.sever.x[i] <- pow(sigma.sever.x[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)
    sigma.sever.x[i] ~ dunif(0,4) #<- sqrt(2)/tau.sever[i] #double exponential errors

    tau.sever.y[i] <- pow(sigma.sever.y[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)
    sigma.sever.y[i] ~ dunif(0,4) #<- sqrt(2)/tau.sever.y[i] #double exponential errors

    tau.sever.z[i] <- pow(sigma.sever.z[i],-2) #~ dgamma(2.5,0.5) #dgamma(1.0E-3,1.0E-3)
    sigma.sever.z[i] ~ dunif(0,4) #<- sqrt(2)/tau.sever.y[i] #double exponential errors
    }

for (i in 1:1){ #L: number of lags
    beta.sever[i] ~ dnorm(0.5,1.0E-3)T(0,)
    zeta.sever[i] ~ dnorm(0.5,1.0E-3)T(0,)

    beta.sever.temp[i] <- beta.sever[i] * step(P.sever-i) #zero if not chosen
    zeta.sever.temp[i] <- zeta.sever[i] * step(Q.sever-i) #zero if not chosen

    beta.cum.sever[i] <-  sum(beta.sever[1:i]) #sum(beta.sever.temp[1:i])
    zeta.cum.sever[i] <-  sum(zeta.sever[1:i]) #sum(zeta.sever.temp[1:i])
    }

for (i in 2:(L+1)){ #L: number of lags
    beta.sever[i] ~ dnorm(0,1.0E-3)
    zeta.sever[i] ~ dnorm(0,1.0E-3)

    beta.sever.temp[i] <- beta.sever[i] * step(P.sever-i) #zero if not chosen
    zeta.sever.temp[i] <- zeta.sever[i] * step(Q.sever-i) #zero if not chosen

    beta.cum.sever[i] <-  sum(beta.sever[1:i]) #sum(beta.sever.temp[1:i])
    zeta.cum.sever[i] <-  sum(zeta.sever[1:i]) #sum(zeta.sever.temp[1:i])
    }

for (i in 1:1){
    chi.sever[i] ~ dnorm(1,1.0E-3)
    psi.sever[i] ~ dnorm(1,1.0E-3)
    nu.sever[i] ~ dnorm(1,1.0E-3)
    xi.sever[i] ~ dnorm(1,1.0E-3)
    }

P.sever <- trunc(P.sever.raw*(L+1)) + 1
Q.sever <- trunc(Q.sever.raw*(L+1)) + 1

P.sever.raw ~ dbeta(1,1)
Q.sever.raw ~ dbeta(1,1)

}
```

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8. REFERENCES

[1] Congressional Budget Office, *Factors Underlying the Growth in Medicare's Spending for Physicians' Services*, **2007**: June, www.cbo.gov/ftpdocs/81xx/doc8193/06-06-medicarespending.pdf.

[2] Hadley, Jack, James Rechovsky, Catherine Corey, and Stephen Zuckerman, "Medicare Fees and Volume of Physicians' Services," *Inquiry*, **2009**: 46(4), 372–390.

[3] Hadley, Jack, and James Rechosky, "Factors Affecting Physicians' Medicare Service Volume: Beneficiaries Treated and Services per Beneficiary," *International Journal of Health Care Finance & Economics*, **2006**: 62(3), 131–150.

[4] Heaton, Paul, "The Impact of Health Care Reform on Workers' Compensation Medical Care: Evidence from Massachusetts," **2012**, www.rand.org/pubs/technical_reports/TR1216.html.

[5] Hill, Peter, "Lowe Indices," **2008**, World Congress on National Accounts and Economic Performance Measures for Nations, Washington DC, May 13–17, www.indexmeasures.com/dc2008/papers/Lowe%20indices%20revised.doc.

[6] Hurvich, Clifford M., and Chih-Ling Tsai, "Regression and Time Series Model Selection in Small Samples," *Biometrika*, **1989**: 76, 297–307.

[7] Hyndman, Rob J., and Yeasmin Khandakar, "Automatic Time Series Forecasting: the Forecast Package for R," *Journal of Statistical Software*, **2008**: 27(3), 1–22, http://www.jstatsoft.org/v27/i03/paper.

[8] International Labour Office, *Consumer Price Index Manual: Theory and Practice*, Geneva (Switzerland): International Labour Office, **2004**.

[9] McGuire, Thomas G., and Mark V. Pauly, "Physician Response to Fee Changes with Multiple Payers," *Journal of Health Economics*, **1991**: 10, 385–410.

[10] Schmid, Frank, and Nathan Lord, "The Impact of Physician Fee Schedule Introductions in Workers Compensation: An Event Study," **2013**, Casualty Actuarial Society Forum, 2013: Spring, 1–36, www.casact.org/pubs/forum/13spforum/SchmidLord.pdf.

[11] Tukey, John W., *Exploratory Data Analysis*, Reading (MA): Addison-Wesley, **1977**.

Abbreviations and Notations

AMA, American Medical Association
CPI, Consumer Price Index
CPT, Current Procedural Terminology
IQR, Inter-Quartile Range
MAR, Maximum Allowable Reimbursement
NCCI, National Council on Compensation Insurance

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