Three Papers on Physician Labour Supply in Canada

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Lay Abstract

Despite the record-high number of physicians per person, Canadians are experiencing shortages of physicians, which are reflected in long wait times for specialist services and inadequate access to primary care. In this thesis, I use supply and demand data to investigate the reasons for shortages.

Analyzing data from 1987 to 2021, I find that although physicians per person grew significantly, the adjusted growth is negative due to physicians' reduced work hours and population aging. Average work hours decreased mainly because of reduced hours by males and an increasing share of females who usually work fewer hours than their male counterparts.

Investigating data during COVID-19, I find that physicians worked reduced hours during the pandemic's first wave and then resumed regular hours. No gender gap is observed in the reduced hours, but a gap is evident across practice settings, with the hours reduction occurring entirely among those practicing in the community.

Abstract

Despite Canada's record-high physician-to-population ratio, persistent wait times for specialist healthcare and insufficient primary healthcare access raise questions. Why does Canada face medical service shortages notwithstanding its high physician count per capita? What factors should be accounted for in physician workforce planning? To address these questions, I analyze Statistics Canada's population estimates and Labour Force Survey (LFS), and the Canadian Institute for Health Information's (CIHI's) physician expenditure and socio-demographic data, from 1987 to 2021. I focus predominantly on the supply side but also consider the demand side.

In the first paper, I show that despite a 35% increase in physicians per capita from 1987 to 2019, the growth rate adjusted for physician labour supply and population aging is negative four percentage points. A 20% reduction in physician work hours from 1987 to 2020 contributes to this decline. These findings underscore the importance of considering factors beyond physician counts.

In the second paper, examining physicians' COVID-19 responses, I find a statistically significant reduction in work hours during the first wave, with a subsequent recovery to the pre-pandemic level. The net reduction was entirely in community settings, with no statistically significant difference between general practitioners/family physicians and specialists. Moreover, no statistically significant gender differences were observed.

In the third paper, I investigate factors contributing to the declining physician work hours using the LFS– a general-purpose survey. As the LFS survey weights are not designed for physician-specific analysis, I apply a generalized method of moments (GMM) weighting technique using CIHI's physician population data that improves estimation quality. This illustrates how the bias and/or precision of general-purpose surveys can be improved in profession-specific analyses. Reduced hours among males, the increased share of females, workforce aging, and an increase in absence rates and lengths are key reasons behind the decline.

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Declaration of Academic Achievement

The material in this dissertation consists of my research with coauthors. I conducted all of the empirical analyses, and wrote the manuscript jointly with my coauthors from 2020 to 2023. The first paper has been published in the Canadian Medical Association Journal (CMAJ). That paper can be found online at: https://www.cmaj.ca/content/195/9/E335/tab-related-content.

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Introduction

The foundation of Canada's healthcare system rests upon providing universal hospital and medical services to all citizens, seeking to ensure an equitable pathway to medical care supported by taxpayers' money. Therefore, every Canadian deserves access to the service when they need it. However, media reports, academic research, and healthcare intervention data indicate that Canadians frequently encounter unexpectedly lengthy waiting periods for medically essential specialist services, compounded by limited access to primary and emergency care (Canadian Institute for Health Information (CIHI), 2022b; Chan, 2002; Islam et al., 2023; Jones, 2022; Kermode-Scott, 1999; Malko & Huckfeldt, 2017; Purdon & Palleja, 2023; Roos et al., 1998; Smart, 2022; Wait Time Alliance, 2015).

Prolonged wait times for specialist services and the shortfall in primary and emergency care not only intensify individuals' suffering but also results in financial setbacks due to missed workdays, diminished national productivity, and an elevated risk of incomplete recovery. To address the challenges posed by these prolonged wait times and insufficient access, the understanding of the underlying causes is essential for policymakers. In this thesis, I present an in-depth analysis, primarily focused on the supply side—specifically, the availability of physician labour—along with a partial exploration of demand-side factors contributing to the persistent healthcare issues in Canada.

In my first paper, I embark on a quest to unravel a perplexing query: Why, even with the highest per capita physician count in recent years and a perceived surplus in the early 1990s when the ratio was lower, have Canadians been facing long wait times for specialist services and insufficient access to primary and emergency care in recent times? This investigation forms the foundation of my research, in which I strive to connect the supply and demand sides to provide an answer. I conduct a comprehensive analysis combining insights from both sides to accomplish this goal.

Beginning with the supply side, I analyze the self-reported work hour data of physicians sourced from Statistics Canada's Labour Fore Survey (LFS) from 1987 to 2020. Accessing these data through McMaster's RDC, I combine this dataset with historical regulatory registry data on physicians from Canadian Institute for Health Information (CIHI), aiming to enhance the quality of labour supply estimations, facilitated by the reweighting methodology proposed by Hellerstein and Imbens. The data combination is essential because the CIHI data lacks information on labour supply and LFS data for physicians is not representative of the underlying population. This approach allows me to estimate the physician labour supply with improved accuracy.

On the demand side, I integrate CIHI's publicly available expenditure data on physicians with Statistics Canada's population estimates, categorized by age and gender. This integration enables me to calculate physician service demand that has been adjusted for both age and gender. By aligning the population's age- and sex-adjusted demand and changing (reduced) work hours by physicians, a pivotal focal point emerges, providing insights for addressing the ongoing crisis. The finding shows that despite a 35% growth in per capita physicians between the years 1987-2019, the adjusted growth during this period is negative four percentage points. After a simple adjustment, this negative growth of per capita physicians stresses the policymakers' need to change their perspective in planning the future physician workforce.

In my second paper, I explore the responses of the physician workforce to the challenges posed by the COVID-19 pandemic. This investigation is motivated by several critical issues:

1. Even before the pandemic, Canada had been grappling with a physician service shortage. Any unfavorable response from physicians could exacerbate this situation, leading to an

accumulation of unmet needs during this period. Addressing this concern requires immediate attention in workforce planning to address post-pandemic demands effectively.

2. The potential for future crises of similar or lesser magnitude underscores the necessity of drawing lessons from the current pandemic to prepare for such scenarios effectively.
3. Media reports and some surveys suggest that healthcare workers, including physicians, might consider leaving the labour force due to the undue burden imposed by COVID-19.

In tackling these issues, timely data analysis is essential to comprehend the situation fully, enabling practical policy recommendations.

In exploring physicians' responses to the COVID-19 pandemic, I analyze the Labour Force Survey (LFS) data from 2016 to 2022. Extending the data to 2016 allows me to capture the secular and seasonal trends in the pre-pandemic context. During the pandemic, I exploit various waves of COVID-19, recognizing their influence on public health policies and physicians' decisions regarding their work. The findings of this study shed light on significant observations. During the initial wave of the COVID-19 pandemic, there was a substantial and statistically significant reduction in the number of hours that physicians worked per week. However, subsequent waves witnessed little to no change compared to pre-pandemic norms. Most of the decline in hours is observed among physicians whose primary practice locations were in the community. At the same time, those who predominantly worked in hospitals experienced no notable change in their average weekly hours.

At the outset, male physicians exhibited a greater reduction in their working hours than their female counterparts, although they started from a higher baseline. Notably, gender differences emerged between family physicians (with females reducing hours more) and specialists (with females experiencing significantly less reduction in hours). Variations were evident across different regions, and physicians aged over 60 reduced their hours more than their younger counterparts. While exploring how the presence of children affected the reduction in hours, the analysis revealed minimal variation across such groups, except for single parents with a child or children. In this particular group, the reduction in hours during the first wave of COVID-19 was notably lower compared to other child status groups.

The data do not indicate an increase in the likelihood of physicians leaving the labour force or taking prolonged breaks from practice. Contrarily, the probability of taking time off decreased during the third wave. Furthermore, this probability exhibited opposing trends in hospital and community settings, with hospitals showing a decrease in absences. Additionally, the duration of absences decreased during the second trough. Female physicians were observed to reduce the length of their absences more than their male counterparts.

My third research paper builds upon the insights gained from the first paper. In the first paper, I find that physicians in Canada have seen a reduction of over 20% in their work hours between 1987 and 2020. This raises the question of what factors have contributed to this reduction. Understanding the answer to this question is vital for shaping policy recommendations and formulating sustainable physician workforce planning to ensure timely and accessible healthcare services.

To identify the sources behind the decrease in work hours, I examine physicians' work hour data derived from the Labour Force Survey (LFS) from 1987 to 2021. I leverage various socio-demographic characteristics that are provided within the dataset. Similar to the approach in the first paper, I combine the LFS data with physician population data from Canadian Institute for Health Information (CIHI) applying the Hellerstein & Imbens (H&I) method, enhancing the precision of labour supply estimates. However, this paper extends the methodological aspect further. I demonstrate that, within the framework of this paper, the weighting method proposed by H&I offers greater flexibility and improved estimates than conventional weighting methods, specifically the raking method implemented by the Stata package ipfraking, as shown in the Monte Carlo simulations.

Moreover, I apply the H&I method in this paper mostly in estimating weighted average hours of work. In estimating the averages, I discover that the H&I method can be implemented in two steps instead of the original one-step process using the Gauss-Newton algorithm. Under this algorithm, both one-step and two-step methods are equivalent. Employing the method in two steps saves time and demands less computational power. This discovery could promote wider usage of the method, which has been sparingly adopted due to its perceived implementation difficulty.

Utilizing the weighting approach mentioned above, I analyze the patterns in labour supply to uncover the underlying factors contributing to the decline, answering the substantive question of the paper. Despite a significant decrease in the proportion of male physicians in the workforce from approximately 80% in 1987 to 56% in 2021, males still dominate the physician workforce. The significant decrease in their weekly working hours by 7 hours during the studied period, coupled with their significant presence in the workforce, profoundly influences the overall decline observed. The increasing representation of female physicians, which grew from approximately 20% in 1987 to 44% in 2021, characterized by fewer working hours than their male counterparts, influences the declining trends roughly in the same magnitude as the reduced hours from male physicians.

Moreover, I identify the aging of the physician workforce as a crucial factor in the declining trend. The study reveals that the percentage of physicians aged 65 and over has risen from about 8% to 15%, constituting a remarkable 90% increase. This group works significantly fewer hours than their prime-age counterparts. Besides, there has been an increase in the rate of absences from work lasting one week or more, rising from about 9% to 11%. The duration of these absences has also grown from around 4 weeks to 7 weeks, further contributing to the overall decline in average work hours.

Factors like marital or immigration status do not substantially influence the trends in average work hours. However, the interaction between gender and parenthood highlights a significant disparity. Male parents increase their work hours compared to male physicians without children, while female parents reduce their work hours relative to female physicians without children. An analysis based on geography exposes noticeable disparities between regions, particularly during the early years when average work hours are high. However, convergence is observed in the later years as average work hours decrease.

Furthermore, I observe indications of cohort effects, where each successive cohort works fewer hours than their predecessors. However, disentangling these cohort effects from year effects remains a complex challenge, given that secular trends and cohort effects occur simultaneously.

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Chapter 1

Physician Workforce Planning in Canada: The Importance of Accounting for Population Aging and Changing Physician Hours of Work

Key Points

- Although Canada's absolute physician-to-population ratio has increased and is at an historic high, reports of physician shortages and inadequate patient access to physicians abound.
- To reconcile these observations, we analyzed workforce data for physicians from 1987 to 2020, and adjusted the population size to address population aging and the number of physicians to account for changing hours of work.
- Although the unadjusted physician-to-population ratio in 2019 was 35% higher than it was in 1987, we found that full adjustment showed the ratio to be about 4% lower.
- The analysis shows that measures of physicians per capita need to be complemented with both demand- and supply-side adjustments to inform planning for health human resources in medicine; relying on simple trends in physicians-per-capita ratios for workforce planning is not helpful.

1.1. Introduction

Canada has long struggled to maintain an appropriately sized physician workforce.[1–4] The recruitment of foreign-trained physicians over recent decades and, starting in the mid-2000s, increased domestic enrollments in medical schools has led to Canada currently having an historically high physician-to-population ratio.[5] However, concerns about physician shortages and burnout,[6–8] as well as limited access to physician care,[9,10] continue.

Previous analyses of physician supply and demand have not adjusted for both population aging and evolving physician hours of work, despite discussions of these factors being quantitatively important.[11,12] To provide insights into the aforementioned challenges — and to inform the profession, the public and governments in planning regarding the appropriate number of new physicians who should enter practice — we analyzed data from 1987 to 2020 to quantify increasing demand because of an aging population and changing service supply given declining physician self-reported hours of work (Box 1).[13–15]

1.1.1. Box 1: Methods

We obtained billing data on physician care by patients' age and sex, and annual counts of practising physicians from the Canadian Institute for Health Information (CIHI). Statistics Canada provided population demographics and the monthly Labour Force Survey microdata. From the survey, we calculated a consistent annual measure of self-reported hours of work by physicians. The annual sample of physicians ranged from about 1700 to 3800 from 1987 to 2020, for a total of about 93 000. As far as we are aware, using the survey to analyze the supply of physician hours is novel. Previous estimates of hours, generated from different surveys by the Canadian Medical Association (CMA) and its partner organizations, used varying methodologies and are available for only a subset of relevant years.[13] The CIHI provides a full-time equivalent

(FTE) measure in its Physicians in Canada report; however, this is derived from clinical payments and does not provide consistent trends over time.[14,15]

To adjust for population aging, for each year after 1987 (the base year), we calculated how much larger a population with the 1987 age–sex distribution would need to be to require the same number of hours of medical care as that later year. This allowed standardized comparisons for age and sex across populations in different years, while holding services per age–sex cell constant. Relative physician hours by age–sex group were proxied by relative physician expenditures because data on hours by patient age and sex do not exist. We held service provision constant at its 1996 level, which is the earliest year with available data.

To generate each year's supply of FTE physicians, we defined 40 hours of self-reported work per week to be 1 FTE. However, for much of our analysis we indexed the results relative to 1987, and any fixed weekly hours would have produced the same index. Details on the data and methodology are provided in Appendix 1 (available at https://www.cmaj.ca/content/cmaj/suppl/2023/02/28/195.9.E335.DC1/221239-ana-1-at.pdf).

1.2. Why is adjusting for population aging and physician hours of work important?

Unadjusted analyses have shown that the physician workforce grew at a much faster rate than the overall population between 1987 and 2020 (93% v. 43% growth).[5,16] This begs the question: If Canada had a perceived physician surplus in the late 1980s, and the size of the physician workforce has subsequently increased at a rate faster than the rate of population growth, why has patient access to care become problematic?[3] Adjusting for changing population demographics and physician hours of work can offer an answer.

Figure 1.1 presents our findings of trends in population numbers and physicians in Canada from 1987 to 2020, both before and after adjustment for population aging and changing physician average weekly hours of work. In panel A, the age–sex-adjusted population quantifies the increasing demand for physician services relative to the unadjusted population. As panel B makes clear, the full-time equivalent (FTE) supply of physicians is lower than the absolute count (except, by construction, in 1987). Adjustment for physician hours of work showed that, starting in the mid-1990s, a gap opened between the number of physicians and the FTE number of physicians. (Caution is required in interpreting 2020, which coincided with the onset of the COVID-19 pandemic.)

1.3. What is the trend in physician hours of work?

Figure 1.2 depicts trends in self-reported physician weekly hours of work over time. These estimates include weeks in which physicians engaged in work and weeks in which they undertook no work (e.g., vacation, illness or caregiving). We show observed (unadjusted) hours as well as similar series adjusted for physician sex (variable as self-reported to Statistics Canada in the Labour Force Survey), and both age and sex. These adjusted series explore the change in average hours of work associated with the increasing share of physicians who self-identify as female and workforce aging. Our observed trends are consistent with those of studies that used more limited data from Canada or comparator nations.[17–21]

In Figure 1.2, the unadjusted trendline (bottom) shows that average hours of work declined by 21.6%, from 49.1 hours per week in 1987 to 38.5 hours per week in 2020. When excluding whole week absences, the decline is 20.6%, from 54.3 to 43.1 hours per week. For comparison, Statistics Canada reported that, for the entire workforce aged 25 years and older, average weekly hours of work in the same years, including absences, declined by 9.5%, from 35.7 to 32.3.[22] Among occupations, physicians have consistently been an outlier with high hours of work.[17]

Age–sex adjustment (the upper line) accounts for 2.5 hours of the 10.6-hour decline seen in 2020 in the unadjusted analysis. Adjusting only for sex, the middle line shows a small gap opening in the early to mid-1990s that does not subsequently expand; the gap is 1.3 hours in 2020. The overall decline in average physician hours of work was only modestly affected by the increasing share of physicians identifying as female. Females had lower average hours, but that average was relatively stable from 1987 to 2020. In contrast, the hours of male physicians declined, and the gender gap in hours diminished across the period. Adjustment to account for the aging of the physician workforce, depicted as the gap between the sex-adjusted line and the uppermost age–sex-adjusted line, begins to show a worrying trend starting around 2005 that is relevant to policy. The gap between the sex-adjusted estimate of hours worked and the age–sex-adjusted estimate represents 1.2 hours per week in 2020.

1.4. How do adjustments for population aging and physician hours of work affect historical trends?

Figure 1.3 shows that, as workforce planners have pointed out, since the late 1980s — when Canada had a perceived surplus of physicians — the absolute number of physicians per capita has only risen over time. Moreover, the per capita supply increased markedly starting in the late 2000s. However, when the needs of Canada's aging population are considered (MDs/age–sex-adjusted 100 000 population in Figure 1.3) much of the increase is seen to be offset by increased demand. By 2020, population aging can be interpreted to have consumed about 18 percentage points of the increase of 34 percentage points in physicians per capita.

Adjusting only for the decline in physician hours of work, the third line from the top of Figure 1.3, generated a larger gap relative to the simplistic “MDs/100 000 population” trend than did the patient age–sex adjustment. Its trajectory also differs. This gap opened earlier, and its rate of increase slowed after about 2005. Acknowledging that physician hours of work recorded in 1987 were unsustainably high, the reduction in hours by 2005 was roughly equivalent to a 10% reduction in the number of practising physicians relative to 1987 and a 15-percentage–point reduction compared with the MDs/100 000 population curve. By 2019 (avoiding 2020, which was affected by the COVID-19 pandemic), the reduction in hours of work meant that the effective labour supply of physicians was 25 percentage points lower than the conventional MDs/100 000 population benchmark.

The bottom curve in Figure 1.3 combines the 2 adjustments and shows — at the lowest point, in 2005 — a 17% reduction in FTE physicians per age–sex-adjusted Canadian compared with 1987. A rebound followed that was driven by the growing number of practising physicians, and, by 2019, the number of FTE physicians per age–sex-adjusted population was 4% below that in 1987, and 39 percentage points below the unadjusted line. Figure 1.3 illustrates the substantial empirical differences in estimates associated with relatively simple adjustments.

1.5. Is there regional variation in physician supply across Canada?

Plots comparable to those in Figure 1.3 are presented in Appendix 2, available at https://www.cmaj.ca/content/cmaj/suppl/2023/02/28/195.9.E335.DC1/221239-ana-2-at.pdf, broken down by region in Canada with the less populous provinces aggregated. Although the fanning out of the 4 curves is broadly similar, the declines in FTE physicians per age–sex-adjusted population before 2005 occurred primarily in British Columbia, Manitoba and Saskatchewan, and Ontario.

Figure 1.4 presents the number of FTE physicians per 100 000 population adjusted for age and sex in Canada and 6 regions employing 3-year moving averages. Substantial heterogeneity is observable across both time and regions. For comparison, the thick line is the Canadian average. Although Figure 1.4 displays hours rather than indexes relative to 1987, the U-shaped pattern observed in the bottom curve of Figure 1.3 is visible for most but not all regions. Alberta and the Atlantic region do not follow the national pattern.

The variation across regions, except for Ontario (an outlier in the early years), was much larger in 2020 than in 1989. Moreover, Ontario experienced one of the largest swings in physician supply, and Manitoba and Saskatchewan have consistently had a relatively low physician supply.

1.6. What does this analysis mean for physician workforce planning in Canada?

Despite the absolute number of physicians increasing at a faster rate than the general population of Canada between 1987 and 2020, reports of physician shortages and excessive wait times for access to health care began to appear in the mid-1990s. Simple adjustments for population aging and physicians' hours of work reconcile these inconsistent observations. As our analysis shows, the increasing needs of an aging population have been empirically important since around 2005, while the supply of physician service hours has simultaneously declined in a manner that is largely unrelated to the evolving age–sex composition of the physician workforce. Thus, relying on simple trends in physicians-per-capita ratios for workforce planning is misleading and not helpful.

Our analysis does not evaluate trends for the greater part of the COVID-19 pandemic, given the current availability of data from the Canadian Institute for Health Information. However, we did observe a notable dip in hours worked by physicians in 2020, which we consider to be attributable to circumstances related to the pandemic. Any analysis of physician work trends during the COVID-19 pandemic would probably require higher frequency data than the annual aggregate employed herein.

Given the data available to us, we can only speculate regarding the reasons for the decline in working hours beyond the findings presented in Figure 1.2. We consider that about a quarter of the decline in hours worked could be associated with an increasing proportion of older and female physicians in the workforce, as these groups worked fewer hours, on average, than the historical norm. Another possibility is that physicians reducing their average number of hours was a response to physician burnout.[6,7,23–25] Burnout may be driven, in part, by the combination of clinical and nonwork responsibilities, with the latter possibly increasing.[19,21] In a 2022 cross-sectional survey involving Ontario physicians, improved work–nonwork balance was cited as 1 of the top 3 potential solutions to burnout.[26] Another possible explanation is that the proposal that compensation increases exceeding target incomes have caused a cut-back in service provision.[27,28]

Although our analysis is informative, it and the 2 simple adjustments we employed have limitations. For example, we held service provision per age–sex group constant at its 1996 level. This was appropriate for the question under evaluation, but it is likely that service provision per standardized person may have increased over time. Moreover, although the Labour Force Survey data (Statistics Canada) had the advantage of being consistently defined, the survey measured only total hours of work and did not allow hours of direct patient care to be separated from hours dedicated to activities such as administration, research and continuing medical education.[17,29] If hours of direct patient care have reduced as a share of total hours, then the FTE adjustment employed in our analysis was too small in later years. Changing technology and practice styles likely also need their own adjustments. Furthermore, we chose not to disaggregate general and family practitioners from other specialists. This extension was feasible with these data but would make this analysis longer while not adding to the basic message.

1.7. Conclusion

Simple adjustments for aging and physician work hours do not answer all questions about the disconnect between the growth in physicians per capita and reports of physician shortages. However, we have shown that such adjustments can have substantial impact for understanding long-term trends in physician workforce availability. Our analysis also helps to reconcile public experiences of physician shortages with the perceptions of those who focus on unadjusted trends that show rising numbers of physicians per capita in Canada. Planning for physician supply should take adjustments such as these into account as a matter of course, especially given the length of training for new physicians.[30] Pursuing a range of additional adjustments in future estimates would be helpful in efforts to maintain an appropriately sized physician workforce in Canada.

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Figures and Tables

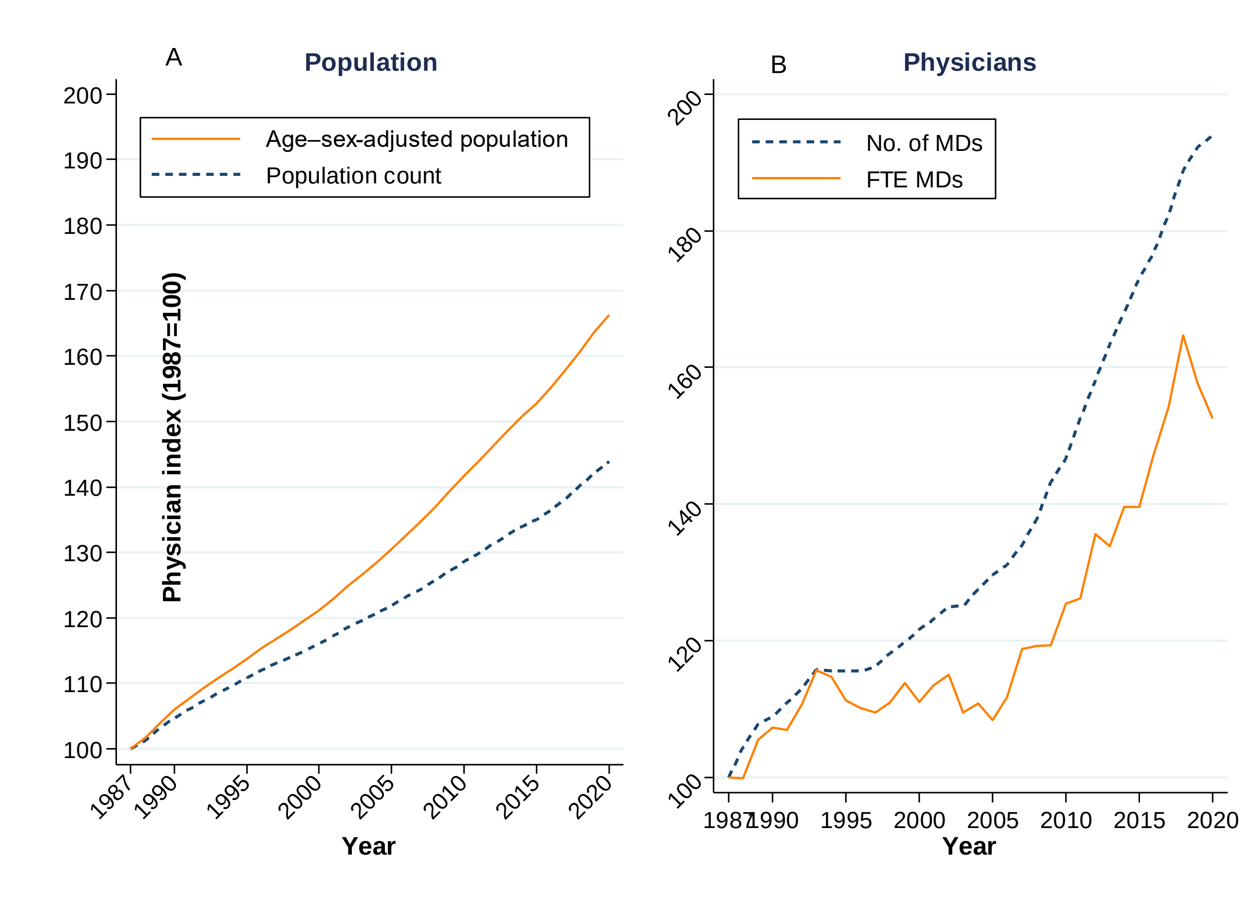

Figure 1.1: (A) Unadjusted and age–sex-adjusted population count, and (B) absolute and FTE physician numbers in Canada from 1987 to 2020.

Note: FTE = full-time equivalent.

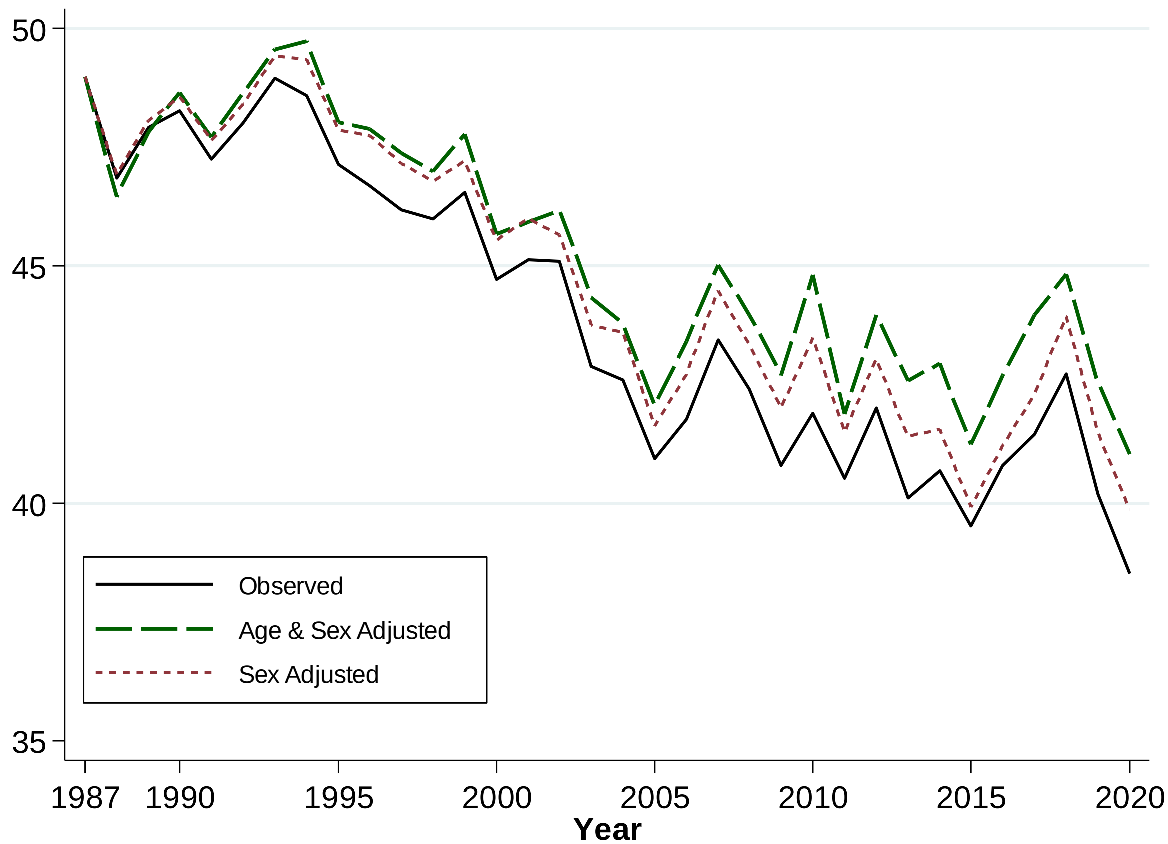

Figure 1.2: Weekly average hours of work by physicians in Canada, including those away for full week, observed unadjusted and adjusted for sex, and age and sex, from 1987 to 2020.

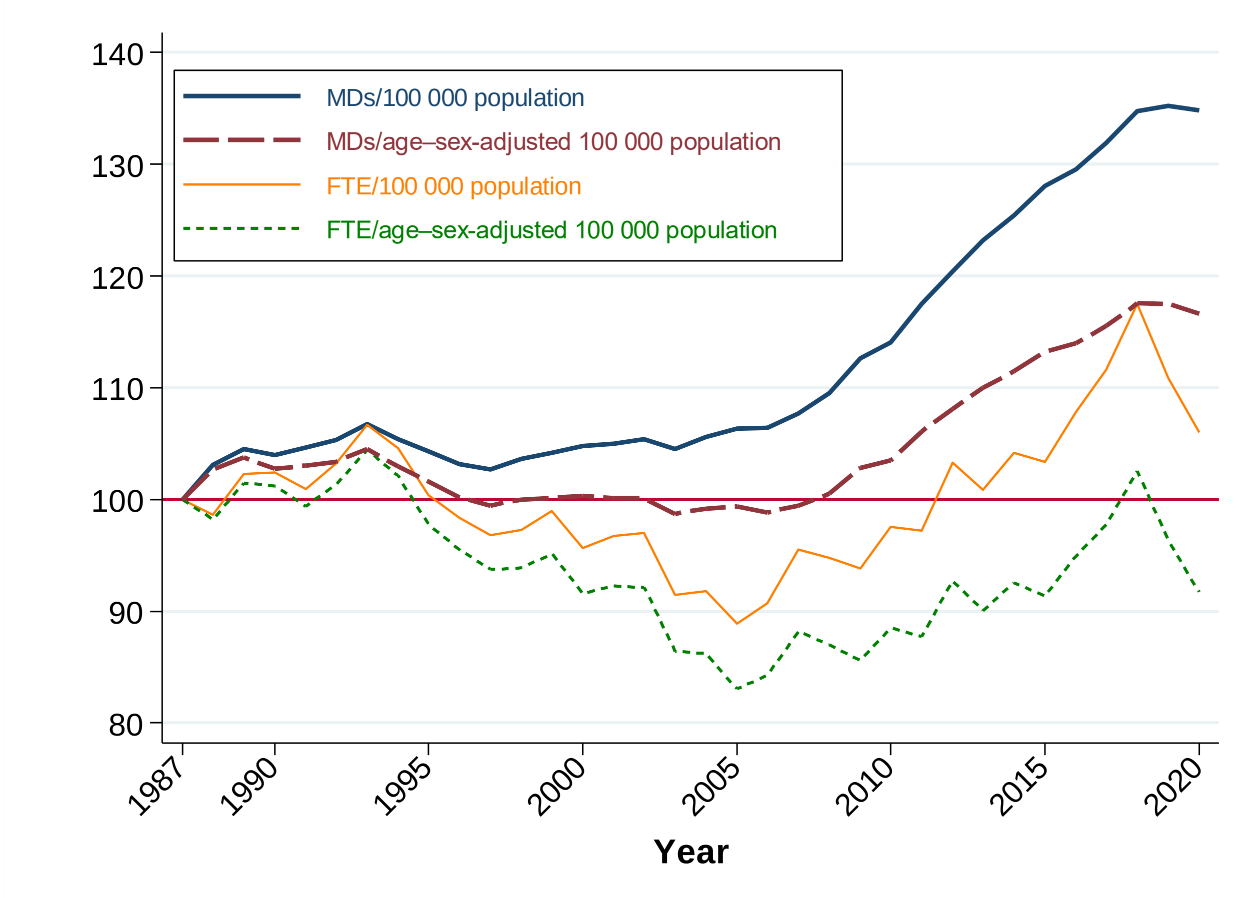

Figure 1.3: Physicians per capita, adjusted and unadjusted (index 1987 = 100, red solid line).

Note: FTE = full-time equivalent.

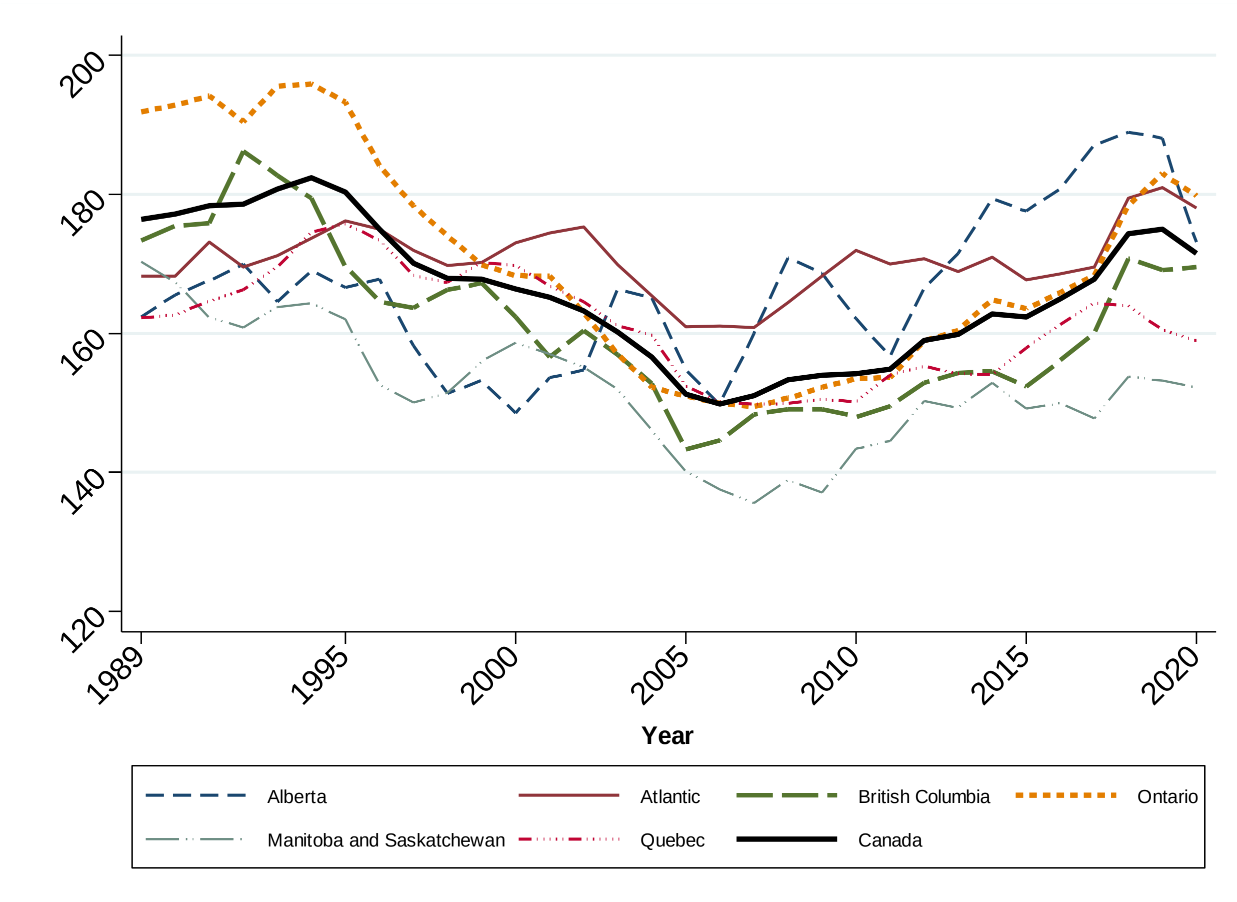

Figure 1.4: Regional distribution of FTE physicians per age–sex-adjusted 100 000 population (3-year moving averages).

Note: FTE = full-time equivalent.

Appendix 1.1: Data and Methods

Data Sources

We combined four different datasets for this analysis. Statistics Canada's confidential monthly Labour Force Survey (LFS) masterfile was accessed through the Statistics Canada Research Data Centre (RDC) at McMaster University. The LFS is a general-purpose survey that is representative of the Canadian population aged 15 and over.[1]

Three public use datasets were also employed: Statistics Canada's Population Estimates,[2] the Canadian Institute for Health Information's (CIHI's) historical data on the physician supply in Canada,[3] and CIHI's data on physician expenditures by age and sex.[4] For the analyses focusing on provinces/regions, we employ province/region-specific population and expenditure measures. CIHI provides a full-time equivalent (FTE) measure in its annual Physicians in Canada report, but it adjusts based on clinical payments. This does not allow measures of trends over time in labour supplied, which are central to this analysis.[5,6]

Data Preparation and Variable Definitions

Monthly LFS files were aggregated into annual files from 1987 to 2020. We restricted the earliest date to 1987 because prior to then relevant occupational codes were unavailable in the most recent version of the LFS masterfile. The latest date is 2020 since the CIHI data were only available until 2020.

The annual samples of physicians grew from about 1700 to about 3800, and the pooled sample for 1987-2020 was around 93,000 (rounded to satisfy Statistics Canada's disclosure rules). We identified physicians using the National Occupational Classification (NOC), version 2016. The occupational code for family physicians and general practitioners is 3112, and that for specialist physicians is 3111. Using the North American Industry Classification System (NAICS), version 2012, we included only respondents whose "main job" was in an industry compatible with a practicing physician. Observations with the following NAICSs codes were included in our sample: 6211, offices of physicians; 6214, out-patient care centers; 6215, medical and diagnostic laboratories; 6216, home health care services; 6219, other ambulatory health care services; 6220, hospitals; 6230, nursing and residential care facilities; 9112, other federal government public administration; 9120, provincial and territorial public administration; 6113, universities; 5417, scientific research and development services; 5241, insurance carriers; and 3254, pharmaceutical and medicine manufacturing.

To match the CIHI administrative data's exclusion of residents as closely as possible, we excluded full-time students and those aged below 28 from our analysis. Therefore, most resident physicians are not included in our sample. We did not impose a restriction on the upper limit of age and rare physicians over age 75 are observed; the oldest practicing physician across all the years in the sample was in the 85 to 89 range. The LFS data employed do not include those residing in the Territories. The LFS supplies general-purpose survey weights, however, since we focused on a small and well-defined sub-population for which administrative data exist that may be used as a benchmark, we improved our estimates' precision by generating new survey weights using CIHI's physician population data within a GMM framework.[7,8] In practice, this made a modest difference and did not change the qualitative interpretation of the results.

We calculated the number of FTE physician positions each year by, first, multiplying the year's average reported weekly hours of work from the LFS with CIHI's total number of practicing physicians in Canada in that year. Second, we divided the product by 40. We consistently assumed 40 hours a week to equal one FTE. Of course, for much of the analysis we focus on an index using 1987 as the base year (i.e., setting 1987 to equal 100), and such an index is not influenced by the number of hours used to define one FTE if that number is used in all years. Similarly, comparisons across provinces were not affected by the choice of 40 to equal one FTE since the same adjustment was used in all cases. The analysis was conducted using Stata version 17.

Measuring Hours of Work Using Statistics Canada's Labour Force Survey

The decline in hours observed in the LFS and reported in the main paper is slightly steeper than that reported by the Canadian Medical Association (CMA) for the selected years where CMA data were available (1997 to 2004, 2007, 2010, 2014, 2017 and 2019), although the levels in much of the period are comparable.[9,10] However, we would not have expected them to have been identical since the two measures are conceptually distinct. The differences result from survey design issues and choices made in the analysis, and may have also been affected by survey non-response and/or survey weighting techniques.

We first address survey design issues combined with choices made in the analysis. To track hours of work, the CMA used a variety of distinct annual survey instruments with surveys administered in many but not all years. Between 1997 and 2003 it used its annual Physician Resource Questionnaire; in 2004, 2007, 2010 and 2014 it used the National Physician Survey, which was a joint effort of the CMA, the College of Family Physicians of Canada, and the Royal College of Physicians and Surgeons; and in 2017 and 2019 it used the CMA's Physician Resource Survey. These were annual retrospective surveys that asked a question such as (from 2017): "EXCLUDING ON-CALL ACTIVITIES, how many HOURS IN AN AVERAGE WEEK do you usually spend on the following activities? Assume each activity is mutually exclusive for reporting purposes (i.e., if an activity spans two categories, please report hours in only one category).” This was followed by 10 categories, including “other”, and then a “total”.

In contrast to the CMA data sources, the LFS was conducted once per month (always in days around the 15th) and asked questions about jobs held in the most recent week.[11,12] The LFS had a range of distinct hours questions. It asked about “usual” weekly hours for both “all jobs” and the “main job”, where main was defined as the job with the highest number of usual weekly hours. It also asked about “actual” hours worked in the survey week, again separately for all jobs and the main job. We employed last week’s actual hours in all jobs, which eliminated issues of recall bias and the annual averaging (with ambiguity in the treatment of holidays and other absences) that is present for the CMA data. We did not employ usual hours; however we note that, on average, over most survey months actual hours worked were less than usual hours.

The LFS also made distinctions between full and partial (relative to the person’s usual activity) weeks of work, and weeks where the respondent was entirely absent from work (and it captured reasons for part- or whole-week absences such as vacation, illness, caring for others, etc.). In our analysis we wanted a comprehensive definition of actual hours supplied, so we used actual hours of work in weeks with some work and also included full week absentees with zero hours of work. As a result, our decline in hours of work reflected both changes in “actual” hours including changes in weeks of partial work, and changes in the likelihood of full weeks with no work. Conceptually, our measure of hours of work was, therefore, different than that employed by the CMA. We believe that our measure is preferable in focusing on actual (including part-weeks) rather than usual hours per week, and in including changes (a small increase) in full weeks away from work. However, the LFS cannot address sub-categories of work as can the CMA’s surveys.

There were also methodological differences across surveys and measures. The CMA's surveys had the advantage of focusing exclusively on physicians. The older Physician Resource Questionnaire targeted a 15% random sample of the physician population. In the late-1990s response rates were over 80%, but they fell to about 30% by 2003.[13] It is not clear how the declining response rate affect the trend observed in the CMA data. Survey weights appear to have been initiated only in 2004 or 2007 with the advent of the National Physician Survey.

In contrast to the CMA data, the LFS provided consistent survey questions, and although it is a general-purpose survey the annual count of observations on physicians across the 12 monthly surveys was comparable to that in the Physician Resource Questionnaire, although smaller than that in the later surveys. (Note that the LFS has a six-month rotating panel design, so the number of observations is not the same as the number of unique respondents. Although we did not exploit this feature in this analysis, it permits variation in hours for individuals over six-month periods to be documented.) The LFS used survey weights to account for its survey design and nonresponse. While the results using the LFS weights provide similar results to those presented, as mentioned we developed survey weights specific to this analysis taking advantage of the population counts of physicians by age, sex and province available from CIHI. This improved the precision of the estimates, and we believe it is a superior methodology even if the basic pattern of results was unaffected.

The LFS is collected under the Statistics Act, and responses can be required of participants. Statistics Canada's survey methodologists undertake ongoing validation of various aspects of the survey and make scheduled adjustments to update the sampling strategy and other survey elements. Numerous reports are available that document these issues.[11,12,14] We are aware of no LFS validations that focus on physicians (since it is a general-purpose survey), however the LFS sampling methodology includes a special stratum that targets high income households.[14]

Adjusting Physician Utilization for Population Aging

The population, adjusted for physician service requirements, was conceptualized to grow over time because of both growth in the population count, and because of population aging since an older population is "equivalent" (in terms of hours of physician services required) to a younger population that is larger. For consistent comparisons, we held hours of physician service provision per age-sex group constant across the study period; evolving utilization by age and sex, and its implications for the size of the required physician workforce, is a potential extension to this simple analysis.

We were not, however, aware of data on hours of service for each patient age-sex group. We therefore employed relative expenditures on physicians for each patient age-sex group as a proxy for relative physician hours for each age-sex group in adjusting for utilization. We calculated base hours of care per age-sex group by multiplying the proportion of total expenditures for each group with total hours in 1987. We used 1996 as our reference year for expenditures since it is the earliest year for which these data exist for age and sex groups and extend this base to both earlier and later years; if these data had existed for 1987, then we would have selected 1987 for this adjustment. To obtain the yearly age-adjusted population in Canada, we multiplied this constant value for hours per age-sex group with the size of each group in each year; for each year, we then summed across groups. Each year's adjusted population measure, therefore, reflected how much larger the 1987 population would need to be to require the same hours of physician care as that year.

Adjusting Physician Hours for the Evolving Sex, and Age and Sex, Composition of the Physician Workforce

For the central human resource/policy question studied, adjusting for sex and/or age was not needed. Documenting the magnitude of the shifting sex and age composition of the workforce in explaining changes in hours was useful, however, for understanding the degree to which the changes in hours could be "explained" by these two factors.

Sex Adjustment

The goal was to create an adjusted series holding the physician male to female ratio constant to contrast with the unadjusted series. We, therefore, counterfactually, kept the sex composition unchanged at its value in the base year, 1987, but allowed the age distribution of each sex, and the hours of work within each age-sex cell, to vary as they did in the unadjusted data.

Focusing on females, but with analogous calculations for males, the method was as follows. We defined the proportion of female physicians in the base year, 1987, as $PropFem87$. In each year we observed the proportion of female physicians in each age group relative to the total number of female physicians. We calculated $SexAdjFem_{AgeGrp,t} = PropFem_{AgeGrp,t} * PropFem87$, and then multiplied this by the average hours observed for that age group in year t to get $SexAdjFemHrs_{AgeGrp,t}$. Finally, we added up all the age groups within each year and obtained a counterfactual annual total. Note that $SexAdjFem_{AgeGrp,t}$ varied from year to year, but within each year the sum of the proportions across age groups summed to the total proportion female in 1987. Thus, the series was counterfactually adjusted to keep the male-female ratio constant while letting the size of the physician workforce, the age distribution of each sex, and both female and male hours by age group, evolve.

Sex and Age Adjustment

As with the sex adjustment, the goal was to create an age and sex adjusted series to contrast with the other series. In this case we, counterfactually, kept both the age composition of each sex unchanged at its value in the base year, 1987, and simultaneously maintained the proportion female (male) at its 1987 value. But, we allowed the hours of work within each age-sex cell to vary as they did in the unadjusted data, and the size of the physician workforce to evolve.

We defined the proportion of male and female physicians in each age group relative to the total physicians of both sexes in 1987 as $PropMale87_{AgeGrp}$, and $PropFem87_{AgeGrp}$. We counterfactually constrained the age-sex composition to remain unchanged at the 1987 level. Then we multiplied $PropMale87_{AgeGrp}$, and $PropFem87_{AgeGrp}$ with their respective hours for each year (e.g., $AgeSexAdjFemHrs_{AgeGrp,t} = PropFem87_{AgeGrp} * FemHrs_{AgeGrp,t}$). Finally, we summed the age-sex groups within each year to obtain annual age-sex-adjusted average hours. The difference between the age-sex-adjusted and sex-adjusted hours provides the age-adjusted hours.

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14. Statistics Canada. *Methodology of the Canadian Labour Force Survey*. Ottawa, Canada, 2017.

Appendix 1.2: Regional Versions of Physicians per capita, Adjusted and Unadjusted

Note: Given the smaller sample sizes in the regions compared to the total sample, the regional results in this appendix are presented as three-month moving averages starting in 1989, with for example 1989 being the average of 1987, 1988 and 1989. The adjustments for each province/region employ province/region-specific age distributions, and province/region-specific costs by age-/group.

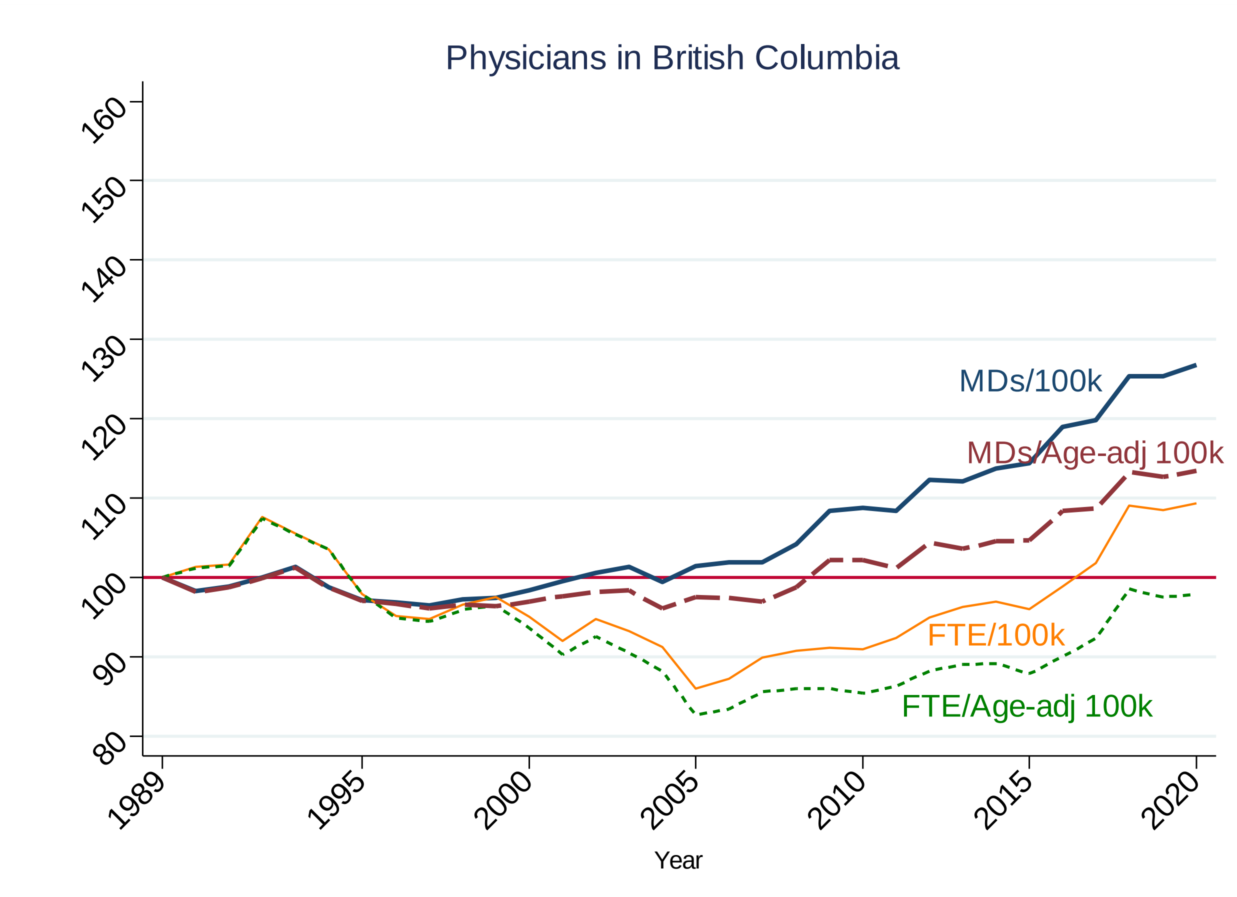

Figure A1.1: Physicians Per Capita in British Columbia, Adjusted and Unadjusted (3 month moving averages)

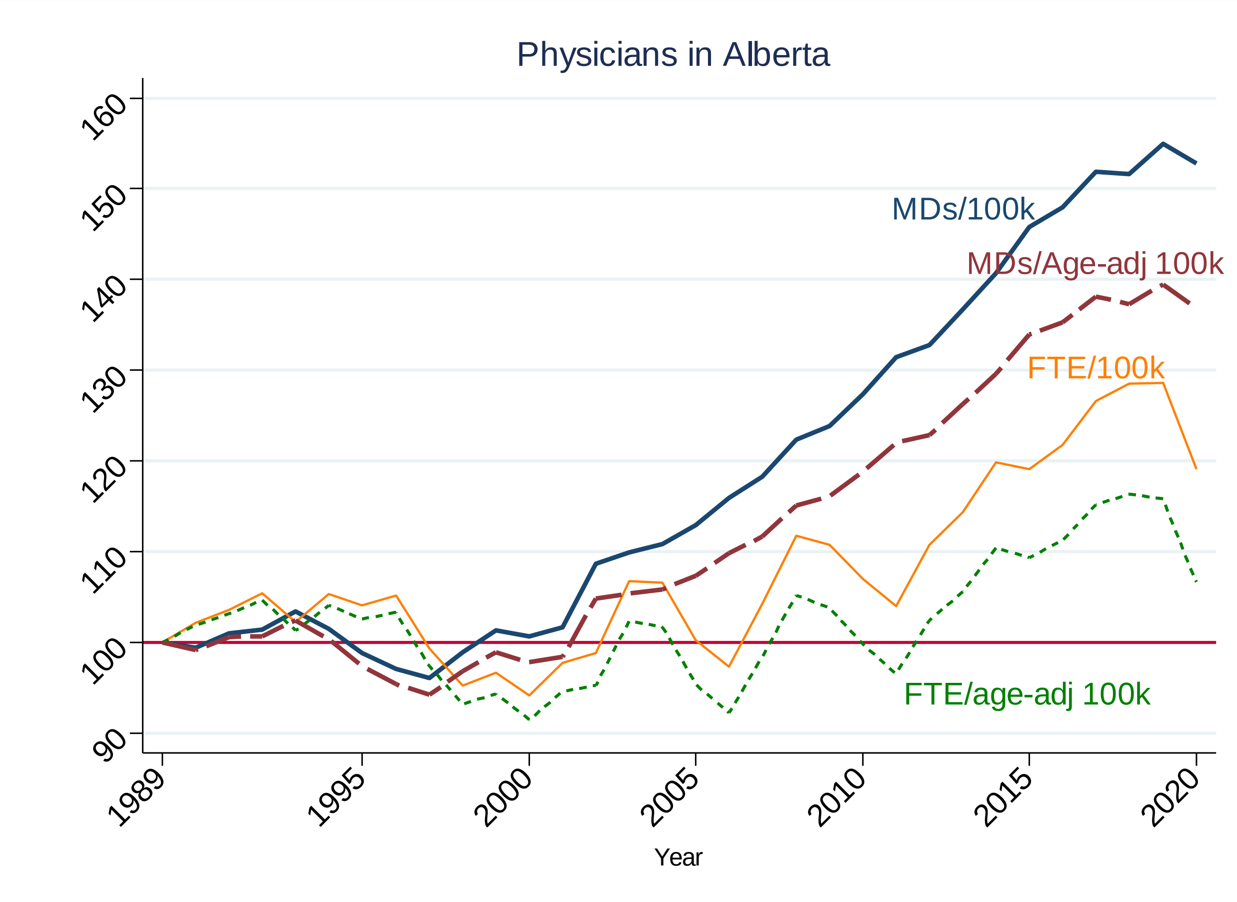

Figure A1.2: Physicians Per Capita in Alberta, Adjusted and Unadjusted (3 month moving averages)

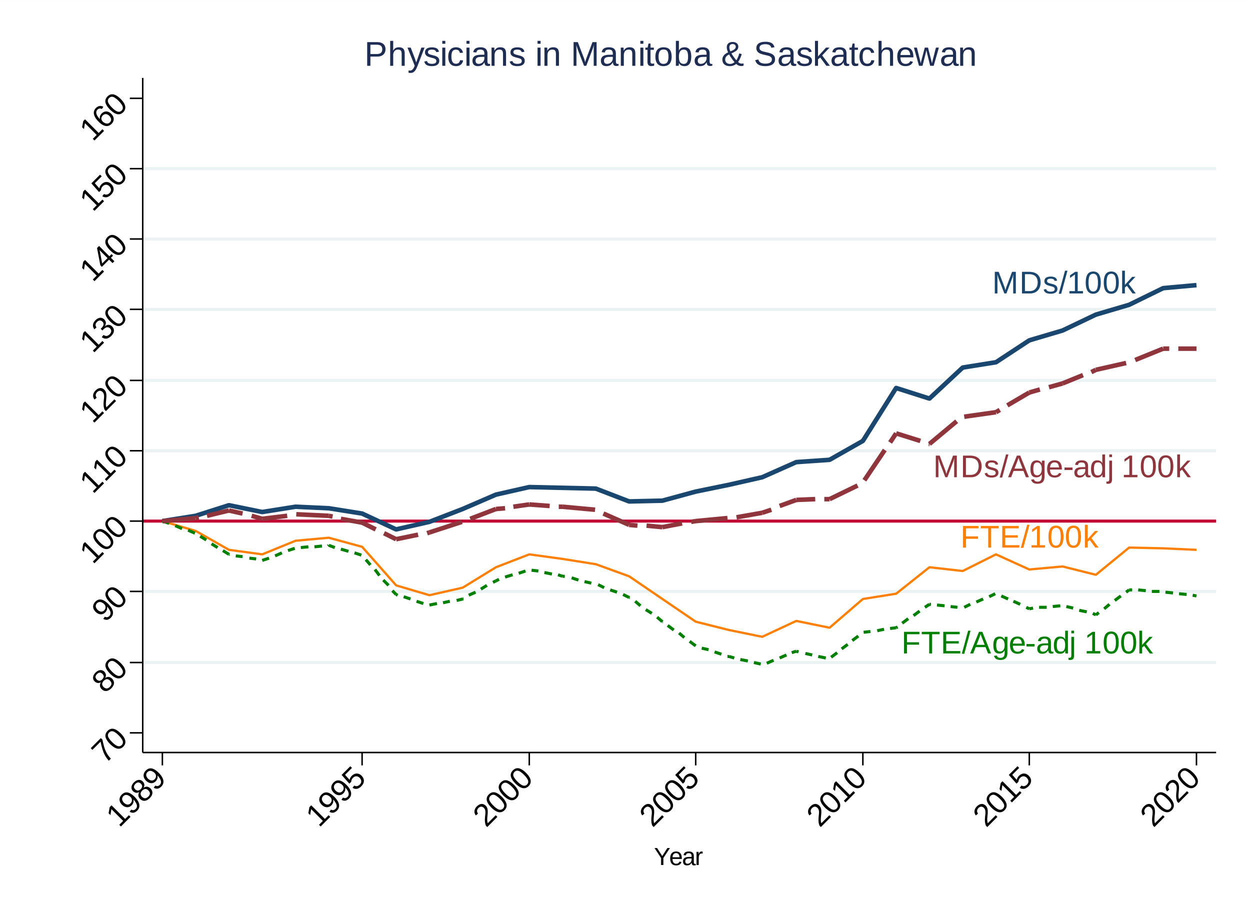

Figure A1.3: Physicians Per Capita in Manitoba and Saskatchewan, Adjusted and Unadjusted (3 month moving averages)

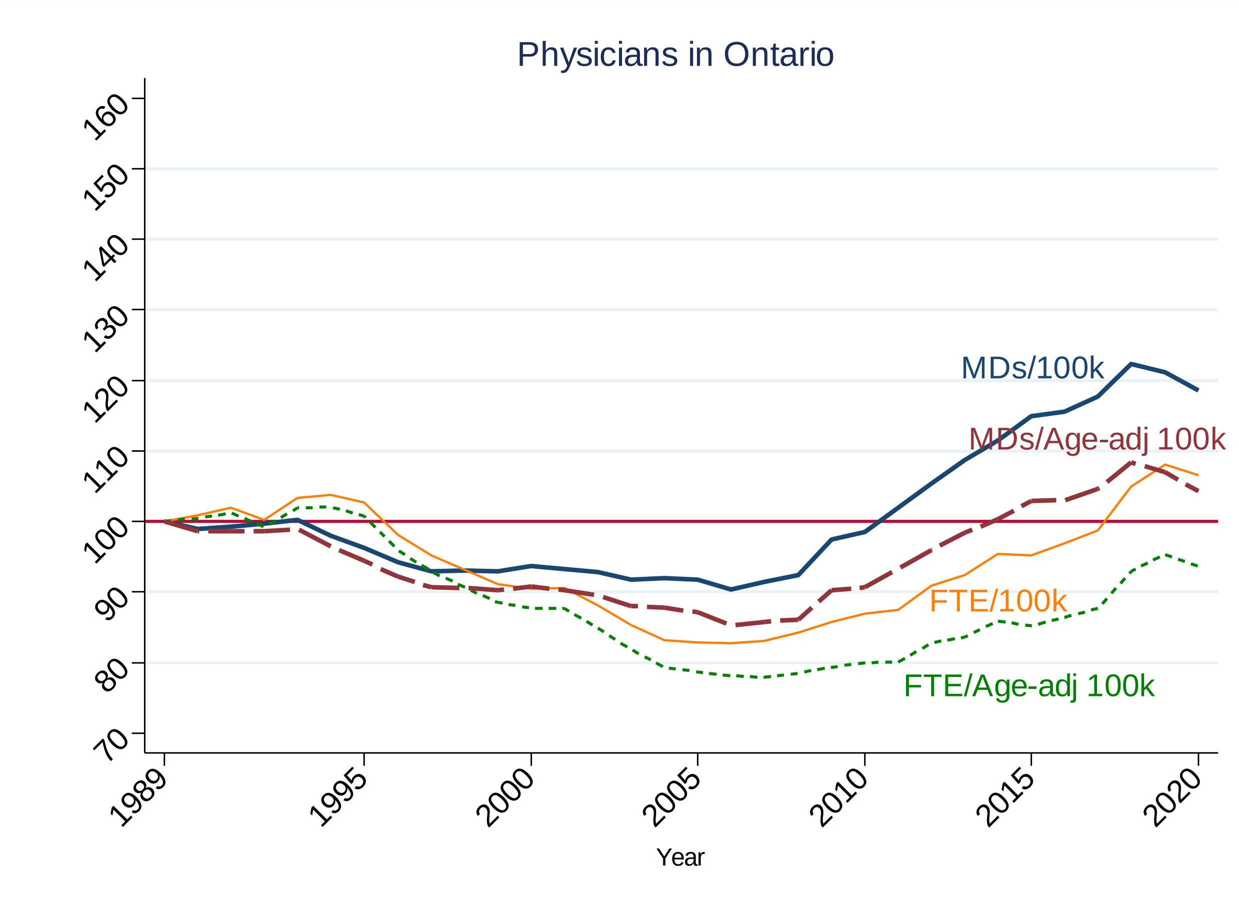

Figure A1.4: Physicians Per Capita in Ontario, Adjusted and Unadjusted (3 month moving averages)

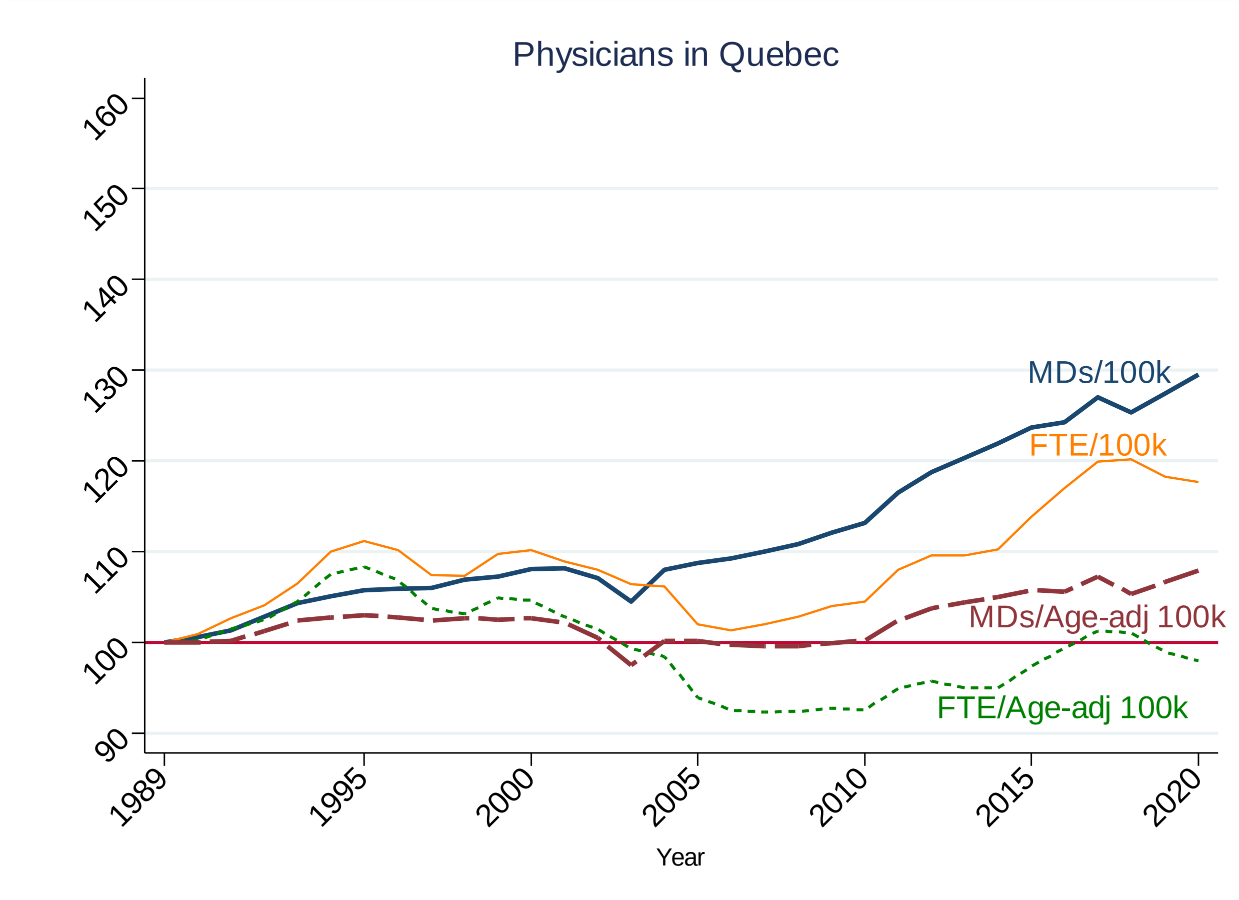

Figure A1.5: Physicians Per Capita in Quebec, Adjusted and Unadjusted (3 month moving averages)

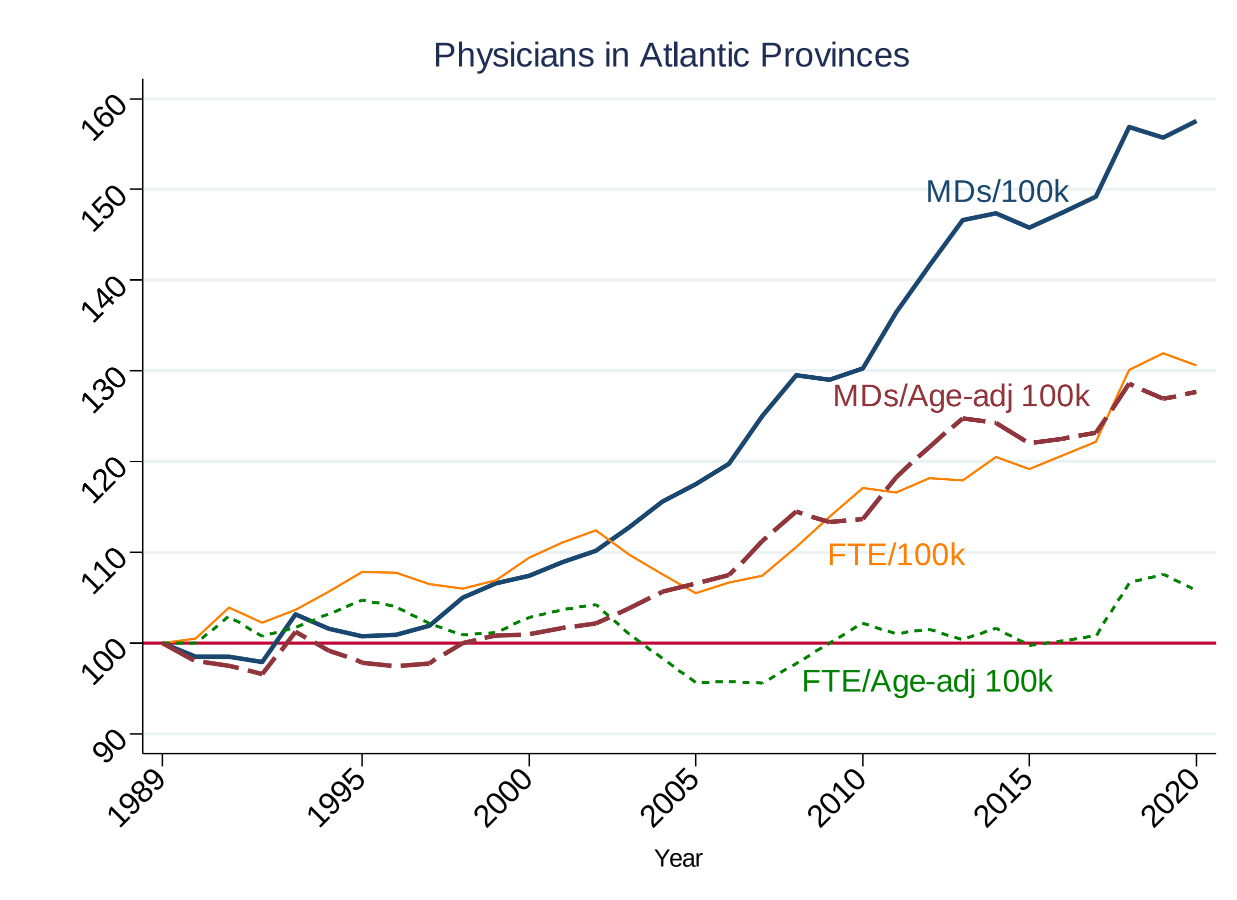

Figure A1.6: Physicians Per Capita in Atlantic Provinces, Adjusted and Unadjusted (3 month moving averages)

Chapter 2

Impacts of COVID-19 on Physician Labour Market Activities

Abstract

Understanding variations in physicians' labour supply during the COVID-19 pandemic can inform health workforce planning. Using nationally representative Canadian data, we observe a reduction in hours in the first wave, but throughout the remainder of the pandemic hours did not differ from pre-pandemic norms. For both family practitioners and other specialists, the hours reduction was concentrated among community-based physicians; hospital-based physicians did not reduce their labour supply. Of note given the large literature looking at differences in male and female physicians' hours of work, both genders had statistically indistinguishable hours changes during COVID-19 with the point estimate for males showing a slightly larger reduction. Further, with the exception of single parents with children whose hours of work were unaffected by the pandemic, marital status and the presence of children does not have appear to have been associated with differences in hours worked. Immigration status similarly did not affect changes in labour supply. However, a larger reduction in hours is observed for physicians over age 60 compared to younger practitioners. Although there were some differences, similar results are observed for absences. No evidence is observed to support the contention that physicians were more likely to exit practice during COVID-19.

2.1 Introduction

Understanding how healthcare workers', and especially physicians', labour supply responded during the COVID-19 crisis is useful for health workforce planning in less extreme contexts as well as for building a robust healthcare system that can weather future crises of various types. In Canada, and other OECD nations, the pandemic occurred in a context with constrained access to primary care and queues for specialist services (Canadian Institute for Health Information (CIHI), 2022; Wait Time Alliance, 2015; Siciliani, 2013). We analyze the labour market activity, especially hours of work and absences from work, of physicians during COVID-19 in Canada.

COVID-19 strained health delivery not only by infecting the general population (e.g., Miller et al., 2020) but by infecting those who treat them (Bandyopadhyay et al., 2020) and also by increasing the threat of infection and inducing stringent infection prevention and control measures (e.g. Centers for Disease Control and Prevention, 2020). Infection and the increased risk of infection among physicians can lead to reduced work hours in weeks with work, increased time away from work, and even earlier than normal retirements. In contrast, professional ethics in the crisis may motivate increased work commitments, postponed vacations and may induce the retired to return to work (Simonds and Sokol, 2009). Government policies such as lockdowns, shutdowns, and limited access to practice settings may also cause physicians to alter their activities (Burdett et al., 2021; Ciminelli and Garcia-Mandicó, 2022), and labour supply may vary with sociodemographic characteristics. By analyzing monthly data from different phases of COVID-19, this study documents changes in the labour market activities of physicians in Canada.

Research in the healthcare sector during COVID has primarily focused on areas other than labour supply, such as shortages of personal protective equipment (PPE) (Eggertson, 2020;

Burki, 2020), service delivery methods (Demeke et al., 2020; Busso et al., 2022), deferring regular physician visits (Cutler, 2021; CIHI, 2021b; Wilensky, 2022), health complications (Zhao et al., 2022; Cannata et al., 2022), infections among physicians (Duong, 2021; Sabetian et al., 2021), and burnout (Restauri and Sheridan, 2020; Khan et al., 2021). Only a few studies (Hu and Dill, 2021; Cutler, 2021; Frogner and Dill, 2022) focus on clinical and labour market activities. The only Canadian report is based on physician billing data, which had major disruptions during the early part of the pandemic and does not address hours of work and absences from work (CIHI, 2021a).

Our data show that during the first wave of the COVID-19 pandemic there was a large and statistically significant decrease in the number of hours physicians worked per week, but subsequent waves saw little or no change relative to pre-pandemic norms. Virtually all of the hours reduction occurred for physicians whose primary practice location was in the community whereas those who primarily worked in hospitals saw no change in average hours worked per week. Of note given the large literature examining gendered differences in physician labour supply (e.g., Ganguli et al., 2020; Wang and Sweetman, 2013), is that male physicians initially reduced their hours more than did females, though from a higher base. However, there were gendered differences between family physicians (females reduces hours more) and specialists (females reduced hours markedly less). Substantial differences are observed across regions, and practitioners over age 60 reduced their hours more than did younger ones. Various approaches to looking at differences in hours reductions as a function of the presence of children show remarkably little variation across such groups, except for single parents with a child/children for whom the hours reduction in the first COVID-19 wave was much less than for other child status groups.

There is no indication of an increase in the probability of leaving the physician labour force or taking a prolonged absence from practice. On the contrary, the probability of taking time off decreased during the third wave. Further, the probability of taking time off varied in opposing directions in hospital and community settings, with hospitals showing a decrease in absences. Additionally, the duration of absences decreases during the second trough. Female physicians reduce the length of absences more than did their male colleagues.

The remainder of the paper is organized as follows. Section 2.2 focuses on the background and data. Section 2.3 explains empirical approaches. Section 2.4 represents and interprets the results. Section 2.5 provides concluding remarks.

2.2 Background and Data

Figure 2.1 illustrates the waves of infection as measured by weekly hospital bed-days due to COVID-19, which are measured more consistently than are infections. A succession of period indicators is defined as the intervals between vertical dashed lines in Figure 2.1 (see Appendix Table A2.1). These are employed in the regression analysis to explore potential changes in physician labor supply. Each period is a full wave (trough to trough) except for: first, the long/low troughs after waves one and three that are identified separately; and, second, the last three waves, which are appended since they are smaller and have higher troughs.

Average weekly hours of work for each successive month are provided by Statistics Canada's Labour Force Survey (LFS). It is a monthly household survey akin to the American Current Population Survey (CPS), which has sometimes been used by U.S. researchers for analyses of the health workforce (Hu and Dill, 2021; Cutler, 2021; Staiger et al., 2010; Frogner and Dill, 2022; Goldman and Barnett, 2023). Besides employment status and hours worked, the LFS also collects information on personal and family characteristics for the civilian, non-institutionalized population aged 15 and older. The LFS has a complex survey design with a panel element, so survey weights and standard errors clustered on the cross-sectional unit (the physician) are employed throughout (Statistics Canada, 2022). The monthly sample size of the LFS is approximately 56,000 households with about 100,000 individuals. We employ files from January 2016, to measure pre-COVID trends and seasonality, to November 2022. The monthly physician sample sizes vary from about 270 to 400. The pooled sample for the entire period is about 28,000 physician-month observations (rounded for confidentiality).

We identify physicians using National Occupational Classification (NOC) codes combined with North American Industry Classification System (NAICS) codes. Only respondents whose "main job" is in an industry compatible with a practicing physician are classified as physicians. (See the appendix for details.) We exclude full-time students and those less than age 28. Therefore, most resident physicians are outside our sample. We do not restrict the upper limit of age, and rare physicians over age 75 are observed; the oldest practicing physician in our data is in the 85 to 89 range. The LFS data do not include those residing in the Territories.

We identify practice setting (i.e., hospital and non-hospital) for the "main job" of each physician; such information is not collected for other jobs (locations of practice). Absences are defined as a week or sequence of weeks in which the physician works zero hours. A part-week absence reduces average hours of work, but is counted as a week of work. Recent labour force exits are identified as individuals who are not in the labour force but who worked within the last 12 months. We have no information on the occupation of the last job for individuals who have been out of the labour force were more than 12 months. The LFS has several measures of hours of work. We focus on self-reported "actual" hours of work, encompassing instances of zero hours, that is, whole week absences from work in all jobs in the survey week (the week in each month containing the 15th), which is distinct from "usual" hours of work. Summary statistics for the sample are provided in Table 2.1.

Table 2.1 presents estimates of the proportion physicians in each of various sociodemographic categories and their average weekly work hours before (January 2016-February 2020), during COVID-19 (March 2020 to November 2022), and in the first wave of the pandemic (March to June 2020). The findings reveal that nearly all subgroups, except those working in hospitals and those who are single/separated/divorced/widowed with at least one child, experienced a noticeable reduction in their work hours during the first wave of the pandemic. However, average work hours during the entire pandemic for which we have data (March 2020-November 2022) were similar to the pre-COVID period. The slight decrease observed during the pandemic, as will be seen, is almost entirely attributable to the first wave. For example, before the COVID-19 outbreak, female physicians typically worked approximately 38.75 hours per week on average; this compares to about 38.14 hours during the pandemic. However, in the first wave of COVID-19, their average work hours experienced a statistically significant drop to about 31.36. In contrast, in the hospital setting the results show the opposite effect with average work hours increasing during the overall COVID-19 pandemic -- and the point estimates suggests in increased slightly more in the first wave than later ones.

2.3 Empirical Approaches

We estimate versions of the following ordinary least squares (OLS) model for hours of work using various subsamples of physicians:

$$Hours_{i,t} = \beta_0 + \delta W_t + \beta D_{i,t} + \gamma M_t + \theta R_i + \varepsilon_{i,t} \qquad (1)$$

where, $Hours_{i,t}$ is the hours of work of physician i in month t, W_t is a vector of indicator variables representing different waves and troughs of COVID-19, $D_{i,t}$ is a vector of demographic controls, M_t is a vector of month indicators to control for seasonality, and R_i is a vector of geographic controls. The error term, ε, is clustered at the physician level given that the rotating panel design implies that physicians may be repeated up to six times. The vector δ provides the effects of COVID-19 on hours of work during each wave and trough and the remaining Greek characters are coefficients to be estimated that conform with the variables they multiply. Models for other dependent variables, such as absence from work, length of absences, and exiting the labour force (LF), are similar to equation (1), but employ either logits or negative binomial (NB) models since the dependent variable is binary or count; details are in the Appendix.

To determine the effects of COVID-19 on work hours, we estimate equation (1) using the entire sample and also subsamples identified by sex and practice setting. Initially, we divide our chronology into pre-COVID, and then the five wave and two trough indicator variables displayed in Figure 2.1. However, we conduct F-tests on the set of wave/trough coefficient estimates beyond Wave 1 for each regression and no test rejects the null hypothesis that the additional waves/troughs are jointly equal to zero. Therefore, to facilitate the presentation and increase precision we compress the specification to "Wave1" and "Wave1+" in regressions beyond the initial set. The latter compressed specifications involve subsamples focused on region, immigration status, age and the like.

We conduct Box-Cox tests to determine whether work hours should be included in the regression in levels or logarithms. The tests support using untransformed data. It appears that even though weekly hours of work are strictly positive, reported hours are sufficiently far from zero and not sufficiently skewed to justify the log transformation. Additionally, we perform joint tests on the whole sample and each subsample to identify whether there were discernable trends in pre-COVID period. We cannot reject the null of no trends over this relatively short time span. Therefore, the period before COVID-19, is reflected in the intercept with no time trend. However, we do find strong evidence of seasonality in reported hours worked and we therefore include a series of month indicator variables omitting January. To determine the order of the polynomial for the continuous variable age, we test down across alternative specifications using the Bayesian information criterion (BIC). The quadratic of "age" gives us the lowest BIC, which is also supported by the t-test on individual coefficients on higher-order polynomials.

2.4 Results

2.4.1 Hours of Work

Figure 2.2 displays physicians' weekly average hours of work, after de-seasonalizing. Details of the adjustment of seasonality are provided in the appendix. In March 2020, at the start of the COVID-19 outbreak in Canada, there was a noticeable decline in the average work hours. This reduction is statistically significant, and large compared to the minimum point of the lower bound of the 95% confidence interval of any pre-pandemic month. However, this reduction only lasted for about four months.

Turning to the regression results, notice that in Table 2.2 and all subsequent tables of regression results we suppress the standard errors and do not present the coefficients for control variables, however these are all available in the appendix. The "All" column of Panel A in Table 2.2 shows that during the first wave of COVID-19, physicians decreased their work by approximately 7.5 hours per week on a pre-COVID base of over 40 hours of work per week. (The intercept indicates that, pre-COVID, a non-immigrant married female physician with no children who was working in the community in Alberta in January worked, on average, 42.6 hours per week; see Appendix Table A2.2.) This reduction is statistically significant at the 1% level. However, no other statistically or economically significant changes are observed in subsequent waves compared to pre-pandemic conditions, except perhaps for a slight reduction in the fifth wave that is at the margin of statistical significance. (Of course, seven t-tests per regression are likely to produce some false positives at the 10 percent level of significance.)

Interestingly, the columns to the right show that the results are almost identical for both genders (the p-value of a two-sided t-test with the null hypothesis that the Wave 1 coefficient on males and females are equal is 0.7579), although on average females worked fewer hours pre-COVID. But, there is an economically large gap across practice settings (p-value=0.0008). Hospital-based physicians maintained their pre-pandemic work hours, possibly due to the increased pressure of treating infected patients (CIHI, 2021b). The entire COVID-19 reduction in physicians' hours of work in Wave 1 occurred among those practicing in the community.

The responses to the pandemic among GP/FPs and specialists are in Panels B1 and B2 of Table 2.2. Although the point estimate for specialists is larger than that for GP/FPs, the point estimates in Wave 1 are not statistically significantly different (p-value = 0.2347). There is, however, a statistically significant difference in Wave 1 hours' changes between specialists working in hospitals and elsewhere (p-value 0.0006). Results from the entire sample and all subsamples with the complete list of variables are available in the appendix. A set of joint hypothesis tests with the null that all the Trough and Wave coefficients beyond Wave 1 are jointly equal to zero could not reject the null hypothesis for each regression in Table 2.2 at conventional levels of significance.

As seen in Table 2.3, Panel A, regional variations in the reactions to COVID-19 in terms of hours reductions are also appreciable. The hours drop in Alberta was largest, followed by Ontario (these two province's coefficients are statistically different from each other at exactly the 10% level). The reduction is moderate in the Atlantic Provinces and British Columbia, but no significant change is observed in Quebec, or Manitoba and Saskatchewan. Again, joint tests on coefficients of waves and troughs beyond Wave 1 show that we cannot reject the null that they are jointly not statistically significant, which is also true in the subsequent samples.

In Panel B of Table 2.3, the Wave 1 coefficients for immigrant and non-immigrant physicians are not statistically significantly different (p-value=0.6321). Our analyses show that the oldest physicians', over age 60, hours response was the largest and the youngest the smallest response (p-value=0.0891).

Table 2.4 explores potential differences by household composition, focusing on the presence of children and marital status. Perhaps surprisingly, physicians with and without children, even young children, show no discernible differences (either economically or statistically) in the change in their hours of the onset of COVID-19. The only group with a statistically significant difference from the others are single parents (those that are either never married, divorced or separated/widowed) and have a child (e.g., p-value=0.0093 in a comparison to those who are married/common-law with no children). This group's average hours seem least affected, indeed unaffected, by the advent of COVID-19.

2.4.2 Absences from Work

Although some might think that the likelihood of physicians being absent from work during the pandemic might increase due to the increased risk of infection, our findings, as depicted in Table 2.5 using logit models, indicate otherwise. On average, physicians did not significantly increase their absences during any of the waves, but instead reduced their likelihood of absences during the third wave, although the patterns vary by gender. Male physicians experienced a slight, and only marginally statistically significant, increase in the likelihood of absences in the first wave, which decreased in the first trough. Female physicians showed no significant change until the third wave, when the likelihood of absences decreased. In the third wave, male physicians also experienced a reduced likelihood of absences. However, the difference between males and females is significant in none of the waves and troughs.

As seem in the first row of Table 2.5 and consistent with the hours results seen previously, for the first wave hospital and non-hospital settings show contrasting results with a decrease in the likelihood of absences in the hospital setting and an increase in the community. However, in the third wave, physicians in the non-hospital setting experience a reduced likelihood of absences, while those in hospitals had stable absence rates. Regarding GP/FPs and specialists, they had similar outcomes for absences but in different waves. Importantly, they experienced an increased likelihood of absences in none of the instances. The difference between hospitals and non-hospitals is significant only in the first wave.

2.4.3 Length of Absences from Work

Table 2.6 studies changes in the length of absences using negative binomial (NB) models among those with absences. We find that physicians appear committed to their work during the pandemic, as the average length of their absences does not significantly increase in any of the COVID-19 waves compared to the pre-COVID era. Instead, there is a reduction of about four weeks in trough 2 (the "All" column of Panel A, Table 2.6). However, there are notable differences between male and female physicians. Female physicians reduce the length of absences by about three weeks against an increase of about two weeks by their male counterparts. The gender differences are significant at the 10% level. Female physicians with absences, on average, also reduce the length of their absence in troughs 1 and 2. Male-female differences are statistically significant.

We find similar outcomes in hospital and non-hospital settings, where the length of absences decreases by about four weeks in trough 2. However, in the hospital setting, the second wave shows a significant increase in absence length at the 10% level; of course, this isolated result may be a type-II error. The difference between hospitals and non-hospitals are not statistically significant in any waves or troughs. GP/FPs and specialists separately show that they cut their length of absences, especially in trough 2. The differences between GP/FPs and specialists are not statistically significant.

2.4.4 Exiting the Labour Force (LF)

Recent media reports (Cimons, 2020; Plater, 2020; Campbell, 2022) and research (Khan et al., 2021) suggest that many physicians are retiring or planning to retire due to the stresses and workload associated with the COVID-19 pandemic. However, the results in Table 2.7 using logit models find no evidence to support the claim that the retirement rate was higher than normal in the period under study. We do not observe an increase in physicians exiting the labor force during COVID-19; this result is consistent across genders. However, we do find differences between hospital and non-hospital settings. In hospitals, the likelihood of physicians exiting the labor force decreases, while in non-hospital settings it increases. These results suggest that physicians remain committed to their profession during this challenging time. However, we cannot estimate the impacts on subgroups because the sample sizes are too small.

2.5 Conclusion

Analyzing self-reported work hours from the LFS, our study examines how Canadian physicians adjusted their work hours during the COVID-19 pandemic. We find that physicians temporarily reduced their work hours during the early months of the pandemic and then returned to their pre-pandemic work hours. However, we observed variations in the reduction among some, but not all, sub-groups. Practice setting is a key predictor of a reduction in hours of work. Of particular note given the large literature looking at differences in male and female physicians' labour supply, male and female physicians responded similarly. Further, marital status and the presence of children does not have appear to have been associated with differences in hours worked, nor was immigration status. However, a somewhat larger reduction in labour supply was observed for physicians over age 60 compared to younger practitioners.

Substantial differences in hours of work reductions during the first wave of COVID-19 were observed as a function of practice setting. Hospital-based physicians did not reduce their labour supply, while community-based ones did so. The reduction in work hours occurred entirely in the non-hospital setting for both GP/FPs and other specialists.

Despite the challenging circumstances, physicians in Canada did not increase their rate of taking leaves of absence, or the length of such absences during the waves of the pandemic studied. Instead, they decreased time away during some pandemic waves. Interestingly, in some instances, female physicians reduced the likelihood and length of absences more than their male counterparts. Moreover, we did not observe an acceleration in physician retirement due to COVID-19.

2.6 References

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Figures and Tables

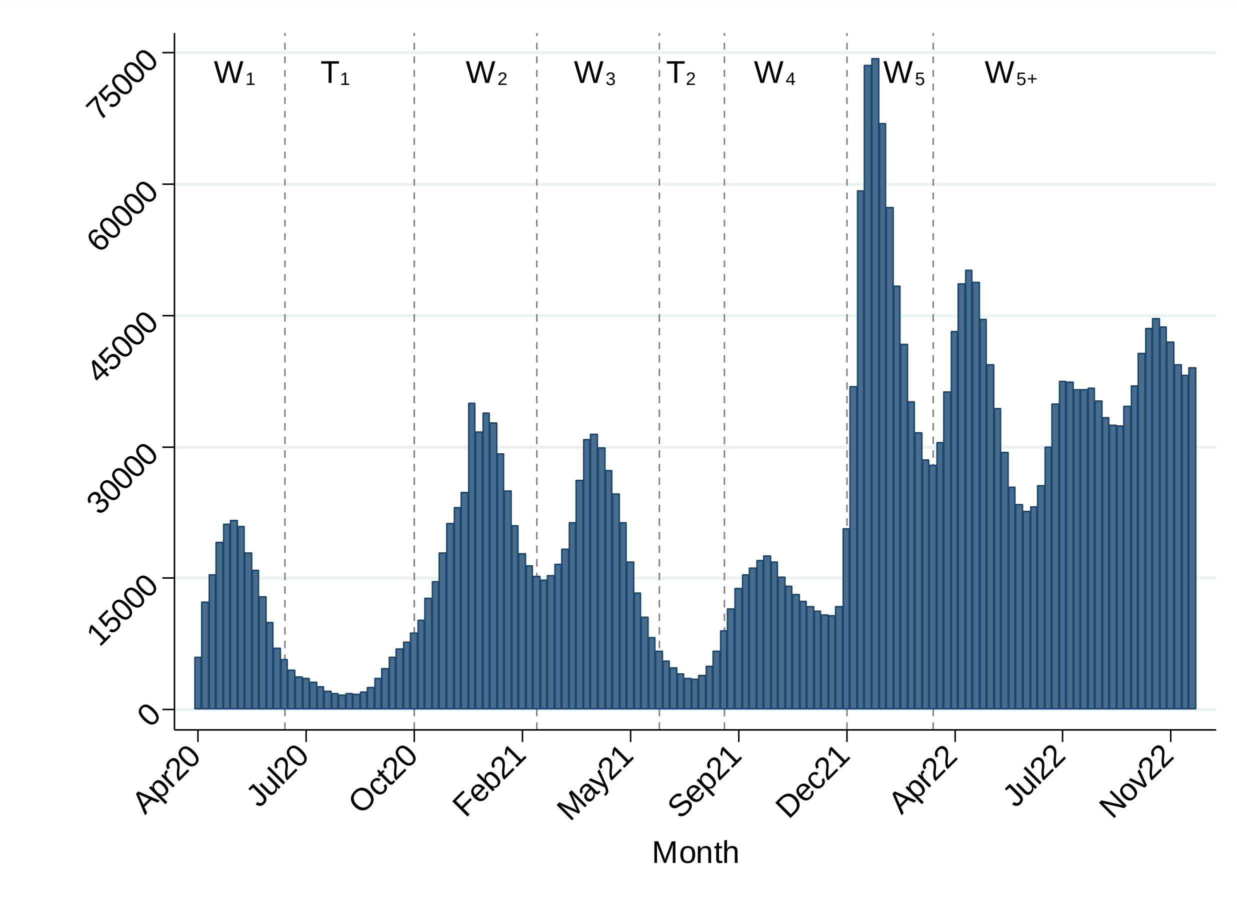

Figure 2.1: Number of hospital bed-days per week occupied by COVID-19 patients.

Source: COVID-19 epidemiology update: Current situation, Health Infobase, Canada; Dec. 19, 2022.

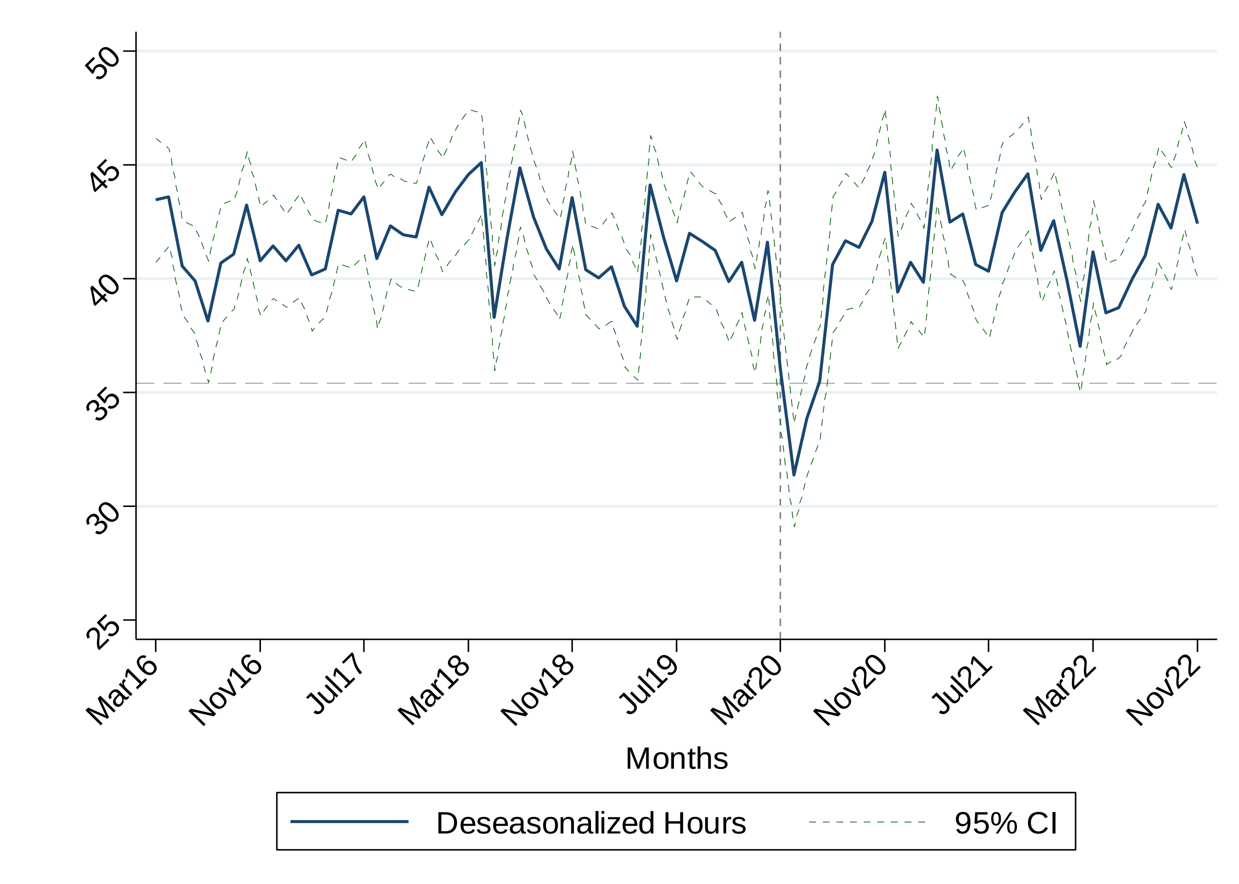

Figure 2.2: Seasonality-adjusted Average Weekly Hours of Work by Physicians (Source: Labour Force Survey)

Table 2.1: The Weighted Proportion of Physicians and Average Weekly Hours of Work

	Proportion (%)	Avg Weekly Hours of Work		
		Pre-COVID (Jan 2016-February 2020)	COVID (March 2020 – November 2022)	Wave1 (March 2020 – June 2020)
Female	46.81	38.75	38.14	31.36
Male	-	43.96	42.64	37.32
GP/FP	59.10	40.78	39.61	34.75
Specialists	-	42.72	41.73	34.96
Immigrant	34.95	41.77	40.53	34.67
Non-immigrants	-	41.51	40.37	34.90
Primarily Hospital	19.49	42.99	43.16	43.92
Primarily Non-hospital	-	41.26	39.78	32.82
Children				
No	47.74	41.59	40.81	35.73
5 Yrs. or Younger	20.30	40.03	38.65	34.98
Only those Aged 6-24	31.96	42.57	41.05	33.46
Children & Marital Status				
Mar/CL W/ Children	43.74	42.01	40.46	33.15
Mar/CL W/O Children	26.08	40.03	40.11	34.02
Other (Single/Sep/Div/Wid) W/ Children	21.66	39.61	38.03	39.80
Other (Single/Sep/Div/Wid) W/O Children	8.52	43.65	41.54	38.00
Region				
BC	14.15	40.42	38.68	34.68
AB	11.39	41.90	38.33	28.82
MN&SK	5.57	41.25	42.32	41.11
ON	43.02	42.25	41.42	34.10
QC	19.31	40.47	39.81	40.55
Atlantic	6.56	43.16	41.96	35.34
Age (Years)				
28-39	33.28	43.06	41.52	38.41
40-49	25.10	42.37	41.62	35.63
50-59	22.57	42.61	41.91	35.23
60 & Above	19.05	37.22	34.64	27.55

Table 2.2: Impacts of COVID-19 on Hours of Work by Sex, and Practice Setting

Dep. Var: Hours of Work	**Panel A:** (All) Physicians				
	All	Male	Female	Hospital	Non-Hospital
Wave1	-7.506***	-7.933***	-7.182***	0.557	-9.323***
Trough1	-0.366	1.923	-3.093*	0.002	-0.370
Wave2	0.101	0.614	-0.291	-3.152	0.729
Wave3	1.422	0.860	2.167	0.294	1.772
Trough2	-0.697	-0.203	-1.147	3.300	-1.391
Wave4	1.095	0.488	1.455	-0.074	1.404
Wave5	-1.994*	-4.278**	0.330	-0.195	-2.297*
Wave5+	-0.023	-0.209	0.278	2.496	-0.521
Intercept	42.608***	48.588***	43.611***	42.683***	42.664***
N	27010	14768	12242	5656	21354
	Panel B1: GP/FPs				
Wave1	-6.281***	-5.236**	-7.732***	-2.520	-7.344***
Trough1	-1.575	1.203	-4.515**	-3.934	-1.295
Wave2	1.161	2.712	-0.289	1.257	1.073
Wave3	2.641	3.086	2.182	3.557	2.643
Trough2	-2.812	-1.822	-3.888	-2.719	-2.681
Wave4	1.257	1.629	0.447	-2.860	2.054
Wave5	-0.994	-1.508	-0.945	4.028	-2.153
Wave5+	-0.862	-0.187	-1.352	0.934	-1.108
Intercept	42.303***	49.222***	42.297***	41.488***	42.523***
N	15973	8419	7554	2678	13295
	Panel B2: Specialists				
Wave1	-9.359***	-11.468***	-6.118**	3.430	-12.628***
Trough1	1.327	3.011	-0.707	2.730	1.064
Wave2	-1.254	-1.844	0.211	-6.146*	0.410
Wave3	-0.439	-2.428	1.745	-2.492	0.307
Trough2	2.836	1.866	5.229	9.783**	1.103
Wave4	1.026	-1.006	4.049	4.067	0.396
Wave5	-3.444	-7.925***	3.248	-6.502*	-2.208
Wave5+	0.968	-0.389	2.472	4.566**	-0.038
Intercept	43.287***	48.070***	45.436***	43.949***	43.164***
N	11037	6349	4688	2978	8059

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, children's status, marital status, immigration status, survey months, and regions. Intercept includes physicians with average age, which is around 48 years old. Full results, including standard errors, are provided in Appendix 2.1.

Table 2.3: Impacts of COVID-19 on Hours of Work by Region, Immigration, and Age

Panel A: Region						
Dep. Var: Hours of Work	AB	Atlantic	BC	ON	MN & SK	QC
Wave1	-14.519***	-7.554***	-7.459***	-8.875***	-1.597	-1.371
Wave1+	0.064	0.553	-0.178	0.131	1.211	-0.286
Intercept	44.554***	40.013***	42.415***	44.636***	43.646***	42.543***
N	2276	4378	3144	8622	3933	4657
Panel B: Immigration and Age						
	Non-immig	Immig	28-39	40-49	50-59	60 & Above
Wave1	-7.207***	-8.399***	-5.592**	-7.407***	-7.460***	-11.105***
Wave1+	0.022	0.164	-0.744	0.801	0.941	-0.897
Intercept	41.099***	43.997***	41.064***	43.983***	38.303***	38.387***
N	18534	8476	8158	7128	6330	5394

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, sex, children's status, marital status, immigration status, survey months, and regions. Full results, including standard errors, are provided in Appendix 2.1.

Table 2.4: Impacts of COVID-19 on Hours of Work by Child, Sex & Child, and Marital & Child Status

Panel A: Child Status				
Dep. Var: Hours of Work	Without Child	With Child	Child <=5 Yr	Child 5+ Yr
Wave1	-7.773***	-7.923***	-6.209**	-8.611***
Wave1+	-0.128	0.172	-0.400	0.517
Intercept	42.372***	40.784***	35.192***	42.701***
N	12744	14266	5478	8788
Panel B: Sex & Child Status				
	Male W/O Child	Female W/O Child	Male W/ Child	Female W/ Child
Wave1	-7.804***	-7.970***	-8.121***	-7.413***
Wave1+	0.473	-0.714	-0.360	0.538
Intercept	48.927***	40.801***	46.575***	42.717***
N	7433	5311	7335	6931
Panel C: Marital & Child Status				
	Married/CL W/O Child	Sing/Div/Sep/Oth. W/O Child	Married/CL W/ Child	Sing/Div/Sep/Oth. W/ Child
Wave1	-8.355***	-6.995**	-9.547***	1.209
Wave1+	1.105	-1.389	0.454	-1.370
Intercept	43.734***	40.391***	40.203***	43.560***
N	7206	5538	11950	2316

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, sex, marital status, immigration status, survey months, and regions. Intercept includes physicians with average age, which is around 48 years old. Full results, including standard errors, are provided in Appendix 2.1.

Table 2.5: Logit estimates of COVID-19 and Absences from Work by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	**Panel A:** (All) Physicians				
	All	Male	Female	Hospital	Non-Hospital
Wave1	0.027	0.044*	0.009	-0.047**	0.044**
Trough1	-0.012	-0.029**	0.013	-0.001	-0.016
Wave2	-0.020	-0.030	-0.012	0.025	-0.029
Wave3	-0.041***	-0.043***	-0.039**	-0.006	-0.047***
Trough2	0.013	-0.004	0.035	-0.001	0.014
Wave4	-0.020	-0.014	-0.025	-0.007	-0.024
Wave5	-0.024*	-0.015	-0.034*	-0.023	-0.024
Wave5+	-0.001	-0.013	0.011	0.020	-0.007
N	27010	14768	12242	5656	21354
Panel B1: GP/FPs					
Wave1	0.028	0.048	0.013	-0.070***	0.046*
Trough1	-0.025	-0.042**	0.000	-0.051	-0.021
Wave2	-0.041**	-0.075***	-0.004	-0.053	-0.040**
Wave3	-0.042***	-0.054***	-0.030	-0.014	-0.045***
Trough2	0.024	-0.008	0.065	0.056	0.021
Wave4	-0.005	-0.021	0.021	0.018	-0.010
Wave5	-0.035**	-0.053**	-0.016	-0.065**	-0.028
Wave5+	-0.010	-0.039***	0.015	-0.004	-0.013
N	15973	8419	7554	2678	13295
Panel B2: Specialists					
Wave1	0.023	0.038	-0.001	-0.031	0.042
Trough1	0.006	-0.016	0.036	0.026	-0.006
Wave2	0.006	0.028	-0.027	0.075	-0.009
Wave3	-0.041**	-0.030	-0.055*	0.009	-0.053***
Trough2	-0.007	-0.002	-0.026	-0.052**	-0.001
Wave4	-0.048***	-0.014	-0.094***	-0.052**	-0.048**
Wave5	-0.006	0.036	-0.063*	0.026	-0.018
Wave5+	0.010	0.020	-0.001	0.033	0.002
N	11037	6349	4688	2978	8059

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, sex, marital status, immigration status, survey months, and regions. Full results, including standard errors, are provided in Appendix 2.1.

Table 2.6: NB Estimates of COVID-19 and Length of Absences by Sex, and Practice Setting

Average Marginal Effect (in week)	**Panel A:** (All) Physicians				
	All	Male	Female	Hospital	Non-Hospital
Wave1	-0.105	2.079*	-3.098*	2.601	-0.652
Trough1	-1.068	0.910	-3.926***	1.628	-0.720
Wave2	2.926	2.110	1.388	11.128*	1.644
Wave3	0.905	-0.291	3.253	12.411	-1.744
Trough2	-3.890***	-1.664*	-6.025***	-3.831**	-3.612***
Wave4	1.545	2.375	-1.442	-2.412	2.300
Wave5	2.400	3.171	-2.095	-1.718	3.368
Wave5+	-0.201	0.041	0.285	2.793	-0.847
N	2704	1295	1409	548	2156
	Panel B1: GP/FPs				
Wave1	-0.319	1.595	-4.264*	3.903	-0.989
Trough1	-1.864	-0.725	-5.055**	7.320	-2.226
Wave2	0.128	1.003	-3.057	20.383*	-1.338
Wave3	-0.340	0.588	-2.082	0.916	-0.720
Trough2	-4.135***	-1.654	-6.615***	-2.069	-4.117***
Wave4	2.113	4.720	-2.373	-2.742	3.360
Wave5	4.314	8.600	-2.150	-2.249	4.830
Wave5+	0.927	-1.552	2.649	3.117	0.201
N	1505	701	804	259	1246
	Panel B2: Specialists				
Wave1	1.317	3.995*	-1.577	-3.940	1.803
Trough1	0.367	2.829*	-2.194	-2.462	2.462
Wave2	9.099	4.294	13.455	15.767	7.778
Wave3	6.116	-0.525	20.630	21.223	-3.207***
Trough2	-2.934***	-0.743	-4.943**	-5.365**	-1.989*
Wave4	-2.170	-0.395	-2.526	0.803	-2.303**
Wave5	1.160	2.113*	-2.817	0.148	1.268
Wave5+	-0.245	2.911*	-3.130	3.106	-0.375
N	1199	594	605	289	910

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, sex, marital status, immigration status, survey months, regions, and year dummies for 2016, 2017, and 2018. Full results, including standard errors, are provided in Appendix 2.1.

Table 2.7: Logit estimates of COVID-19 and Exiting the LF by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	All Physicians	Male	Female	Hospital	Non-Hospital
Wave1	0.016	0.019	0.009	-0.033**	0.026*
Wave1+	-0.003	-0.005	-0.002	-0.023**	0.002
N	27741	15194	12547	5864	21877

*** p<.01", "** p<.05", "* p<.1. N reflects the unweighted sample size.

Note: Controls include age, age squared, sex, marital status, immigration status, survey months, regions, and year dummies for 2016, 2017, and 2018. Full results, including standard errors, are provided in Appendix 2.1.

Appendix 2.1

Notes on Background and Data

Table A2.1: Details Timeline of Various Waves and Troughs.

Wave/Trough	Start Date	End Date
Wave1	March, 2020	June, 2020
Trough1	July, 2020	October, 2020
Wave2	November, 2020	February, 2021
Wave3	March, 2021	June, 2021
Trough2	July, 2021	August, 2021
Wave4	September, 2021	December, 2021
Wave5	January, 2022	March, 2022
Wave5+	April, 2022	November, 2022

Details of NOCs and NAICSs Used in Identifying Physicians

Our samples include observations from the following NOCs and NAICSs. Family physicians and general practitioners are represented by the NOC 3112, while specialist physicians are represented by the NOC 3111. The NAICSs are as follows: 6211 for offices of physicians, 6214 for out-patient care centers, 6215 for medical and diagnostic laboratories, 6216 for home health care services, 6219 for other ambulatory health care services, 6220 for hospitals, 6230 for nursing and residential care facilities, 9112 for other federal government public administration, 9120 for provincial and territorial public administration, 6113 for universities, 5417 for scientific research and development services, 5241 for insurance carriers, and 3254 for pharmaceutical and medicine manufacturing.

Notes on the Empirical Approaches

Models For Other Dependent Variables

We use logit models to estimate the effects of COVID-19 on the likelihood of absences and exiting the labour force (LF). However, models slightly differ as we use all wave and trough dummies for absences and "Wave1" and "Wave1+," including pre-COVID year dummies for the LF exit. We include pre-COVID year dummies in the latter model because joint tests on them support it, which is not true for the model of absences. The following equations represent our logistic model:

$$Pr(y = 1|z) = F(z) \qquad (A1)$$

where $F(z) = \frac{e^z}{1+e^z}$ is the cumulative logistic distribution,

$$z_{i,t} = \beta_0 + \delta W_t + \beta D_{i,t} + \gamma M_t + \theta R_i \qquad (A2)$$ for the model of absences, and

$$z_{i,t} = \beta_0 + \delta W_t + \beta D_{i,t} + \gamma M_t + \theta R_i + \psi X \qquad (A3)$$ for the model of exiting the LF

y can be either absence from work or exiting the LF depending on the model we estimate, W_t is a vector of indicator variables representing different waves and troughs of COVID-19, $D_{i,t}$ is a vector of demographic controls, M_t is a vector of month indicators to control for seasonality, R_i is a vector of geographic controls. Equation (A3) includes ψX, where X is a set of indicator variables for 2016, 2017, and 2018. Further, in equation (A3), W_t includes only “Wave1” and “Wave1+” instead of all waves and troughs that in equation (A2). The vector δ provides the effects of COVID-19 on absences and the LF exit of work during each wave and trough and the remaining Greek characters are coefficients to be estimated that conform with the variables they multiply.

We apply the negative binomial (NB) regression model to estimate the impacts of COVID-19 on the length of absences. The nature of the data leads us to use NB regression. The length of absences is reported in the number of weeks, which is nonnegative. In this section, we consider only the second part of the estimation of the two-part model because we estimate the first part while we deal with the likelihood of absences. Therefore, given that the physicians are absent from work in the first part, the second part utilizes only the nonnegative length of absences to identify any impact on it. We wisorize the data from the upper bound, in which we replace the number of weeks absent from work as 78 if it is longer than 78 weeks- that is a year and half. Given the data, we apply the following equations to capture the impacts of COVID-19 on the length of absences:

$$y_i \sim Poission(\mu_i) \quad (A4)$$

where

$$\mu_i = e^{(\beta_0 + \delta W_t + \beta D_{i,t} + \gamma M_t + \theta R_i + offset_i + V_i)} \quad (A5)$$

And

$$e^{V_i} \sim Gamma\left(\frac{1}{\alpha}, \alpha\right) \quad (A6)$$

Covariates are the same as explained in equation (A2). Here α is the overdispersion parameter that governs variance. In our estimation use dispersion equal to 1+ $\alpha\, e^{(\beta_0 + \delta W_t + \beta D_{i,t} + \gamma M_t + \theta R_i + offset_i)}$.

Notes on Deseasonalizaion in Figure 2.2

We adjust the seasonality in our data extending the pre-COVID timeline to January 2016 following the method proposed by Suits (1984). The adjustment is required because the raw data show noticeable fluctuations in the various months. We obtain the following coefficients by running a regression using work hours data on month indicator variables for the periods of January 2016 and December 2019, i.e., for the pre-COVID years.

$$Hours_{i,t} = \beta_0 + \gamma M_t + \varepsilon_{i,t} \qquad (A7)$$

where M_t represents a vector of month dummies starting from February to December. January is in the intercept, which means it has a coefficient equal to zero. Using the estimated values, we calculate $k = \frac{1}{12}[\gamma_{Feb} + \cdots \dots \dots \dots \dots \dots . +\gamma_{Dec}]$. Then, we project adjusted hours of work for January as $AdjHours_{jan,i} = Hours_{\{jan,i)} - k$ and for other months $AdjHours_{m,i} = Hours_{\{m,i)} - k - \gamma_m$, where $m = Feb, \dots \dots, Dec$. For details, please see Suits (1984).

References:

Suits DB. 1984. Dummy Variables: Mechanics V. Interpretation. *The Review of Economics and Statistics* **66**(1): 177–180.

Detailed Results in Full Models

Table A2.2: Impacts of COVID-19 on Hours of Work of All Physicians by Sex, and Practice Setting

Dep. Var: Hours of Work	All	Male	Female	Hospital	Non-Hospital
Wave1	-7.506***	-7.933***	-7.182***	0.557	-9.323***
	(1.259)	(1.848)	(1.592)	(2.583)	(1.411)
Trough1	-0.366	1.923	-3.093*	0.002	-0.370
	(1.222)	(1.639)	(1.772)	(2.422)	(1.387)
Wave2	0.101	0.614	-0.291	-3.152	0.729
	(1.400)	(2.198)	(1.678)	(2.723)	(1.545)
Wave3	1.422	0.860	2.167	0.294	1.772
	(1.187)	(1.761)	(1.593)	(2.905)	(1.278)
Trough2	-0.697	-0.203	-1.147	3.300	-1.391
	(1.728)	(2.257)	(2.662)	(3.537)	(1.898)
Wave4	1.095	0.488	1.455	-0.074	1.404
	(1.090)	(1.484)	(1.592)	(2.111)	(1.245)
Wave5	-1.994*	-4.278**	0.330	-0.195	-2.297*
	(1.160)	(1.663)	(1.636)	(2.227)	(1.327)
Wave5+	-0.023	-0.209	0.278	2.496	-0.521
	(0.908)	(1.342)	(1.201)	(2.407)	(0.949)
Age	-0.212***	-0.202***	-0.232***	-0.261***	-0.164***
	(0.028)	(0.040)	(0.040)	(0.069)	(0.034)
Age # Age	-0.015***	-0.013***	-0.018***	-0.007	-0.018***
	(0.002)	(0.003)	(0.003)	(0.007)	(0.003)
Male	6.548***			7.801***	6.189***
	(0.573)			(1.141)	(0.643)
Hospital	1.272*	2.106**	0.242		
	(0.696)	(1.074)	(0.875)		
Child<=5	-5.209***	-2.813**	-7.854***	-7.353***	-4.159***
	(0.865)	(1.164)	(1.250)	(1.613)	(1.000)
Child>5	-1.525**	-0.743	-2.854***	-2.555*	-1.254
	(0.740)	(1.009)	(1.082)	(1.506)	(0.819)
Single/Div/Sep/Widow	-0.219	-0.864	0.270	0.092	-0.388
	(0.696)	(1.006)	(0.946)	(1.369)	(0.791)
Immig	0.094	0.389	-0.503	-1.583	0.553
	(0.601)	(0.824)	(0.872)	(1.195)	(0.676)
Feb	-1.275*	-0.763	-1.797*	-0.144	-1.469*
	(0.693)	(0.965)	(0.998)	(1.685)	(0.768)
Mar	-2.309***	-1.433	-3.175***	-1.449	-2.425***
	(0.811)	(1.104)	(1.198)	(1.755)	(0.914)
Apr	-1.893**	-1.823*	-1.842	-1.387	-1.979**
	(0.802)	(1.093)	(1.187)	(1.823)	(0.894)

May	-1.639*	-1.799	-1.251	-3.519*	-1.113
	(0.837)	(1.127)	(1.235)	(1.994)	(0.918)
Jun	-0.753	-1.671	0.379	-0.937	-0.565
	(0.898)	(1.225)	(1.308)	(2.008)	(0.995)
Jul	-4.148***	-5.071***	-3.063**	-3.723*	-4.145***
	(0.943)	(1.270)	(1.380)	(2.093)	(1.055)
Aug	-5.976***	-7.236***	-4.550***	-2.733	-6.768***
	(0.959)	(1.289)	(1.413)	(1.972)	(1.094)
Sept	-0.988	-2.478**	0.796	-0.245	-1.115
	(0.858)	(1.182)	(1.230)	(2.018)	(0.949)
Oct	-5.382***	-6.664***	-3.745***	-3.930**	-5.625***
	(0.822)	(1.110)	(1.202)	(1.996)	(0.906)
Nov	-2.157***	-2.422**	-1.743	1.885	-3.036***
	(0.781)	(1.108)	(1.096)	(1.972)	(0.843)
Dec	0.196	-0.057	0.582	1.918	-0.161
	(0.710)	(0.969)	(1.046)	(1.770)	(0.769)
Atlantic	2.236**	3.290**	1.004	-0.633	3.372***
	(1.043)	(1.398)	(1.562)	(1.780)	(1.243)
BC	-0.187	0.572	-1.262	-2.156	0.154
	(1.095)	(1.440)	(1.665)	(2.422)	(1.223)
Ontario	2.064**	1.937	2.228	0.308	2.399**
	(0.962)	(1.303)	(1.422)	(1.627)	(1.132)
Prairies	1.028	0.166	2.160	-1.915	2.113*
	(1.025)	(1.367)	(1.542)	(1.713)	(1.226)
Quebec	0.652	-0.012	1.427	-0.186	1.002
	(1.105)	(1.520)	(1.596)	(2.486)	(1.228)
Intercept	42.608***	48.588***	43.611***	42.683***	42.664***
	(1.278)	(1.675)	(1.813)	(2.356)	(1.462)
N	27010	14768	12242	5656	21354

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.3: Impacts of COVID-19 on Hours of Work of GP/FPs by Sex, and Practice Setting

Dep. Var: Hours of Work	All	Male	Female	Hospital	Non-Hospital
Wave1	-6.281***	-5.236**	-7.732***	-2.520	-7.344***
	(1.584)	(2.373)	(1.914)	(2.913)	(1.824)
Trough1	-1.575	1.203	-4.515**	-3.934	-1.295
	(1.419)	(1.994)	(2.007)	(3.360)	(1.542)
Wave2	1.161	2.712	-0.289	1.257	1.073
	(1.746)	(2.970)	(1.918)	(3.727)	(1.856)
Wave3	2.641	3.086	2.182	3.557	2.643
	(1.622)	(2.525)	(2.012)	(4.904)	(1.658)
Trough2	-2.812	-1.822	-3.888	-2.719	-2.681
	(2.136)	(2.888)	(3.150)	(5.021)	(2.267)
Wave4	1.257	1.629	0.447	-2.860	2.054
	(1.394)	(2.030)	(1.933)	(2.922)	(1.551)
Wave5	-0.994	-1.508	-0.945	4.028	-2.153
	(1.333)	(2.013)	(1.782)	(2.776)	(1.479)
Wave5+	-0.862	-0.187	-1.352	0.934	-1.108
	(1.202)	(1.822)	(1.500)	(3.914)	(1.161)
Age	-0.162***	-0.143***	-0.205***	-0.292*	-0.109***
	(0.035)	(0.052)	(0.047)	(0.164)	(0.039)
Age # Age	-0.014***	-0.013***	-0.017***	-0.010	-0.016***
	(0.003)	(0.004)	(0.004)	(0.015)	(0.003)
Male	6.634***			5.853***	6.710***
	(0.716)			(1.816)	(0.772)
Hospital	1.522	1.190	1.484		
	(1.057)	(1.774)	(1.240)		
Child<=5	-5.648***	-4.100***	-7.487***	-7.913***	-4.515***
	(1.106)	(1.515)	(1.552)	(2.515)	(1.237)
Child>5	-2.267**	-1.709	-3.257**	-4.488*	-1.950*
	(0.934)	(1.346)	(1.285)	(2.366)	(0.999)
Single/Div/Sep/Widow	-0.025	-1.209	1.066	1.773	-0.451
	(0.857)	(1.314)	(1.079)	(2.069)	(0.914)
Immig	-0.145	-0.462	0.001	-1.747	0.256
	(0.757)	(1.104)	(1.022)	(1.829)	(0.819)
Feb	-1.552*	-0.548	-2.517**	-3.169	-1.179
	(0.882)	(1.270)	(1.215)	(2.492)	(0.942)
Mar	-2.635**	-1.521	-3.664**	-1.931	-2.565**
	(1.031)	(1.445)	(1.461)	(2.342)	(1.147)
Apr	-3.099***	-2.325	-3.868***	-0.624	-3.526***
	(1.021)	(1.431)	(1.453)	(2.367)	(1.115)
May	-2.093*	-0.852	-3.366**	-2.318	-1.993*
	(1.083)	(1.512)	(1.521)	(2.569)	(1.174)
Jun	-1.301	-1.509	-1.053	0.183	-1.445
	(1.123)	(1.622)	(1.528)	(2.738)	(1.209)

Jul	-4.404***	-4.266**	-4.686***	-1.810	-4.878***
	(1.168)	(1.707)	(1.572)	(2.934)	(1.265)
Aug	-4.834***	-5.135***	-4.551***	0.969	-6.003***
	(1.215)	(1.712)	(1.692)	(2.863)	(1.337)
Sept	-1.721	-3.000*	-0.398	1.787	-2.355**
	(1.097)	(1.619)	(1.453)	(3.100)	(1.166)
Oct	-4.650***	-5.588***	-3.639**	-2.831	-4.870***
	(1.059)	(1.501)	(1.464)	(3.096)	(1.118)
Nov	-2.393**	-1.893	-2.814**	3.324	-3.270***
	(0.990)	(1.503)	(1.280)	(3.117)	(1.036)
Dec	0.630	0.862	0.340	2.599	0.367
	(0.904)	(1.325)	(1.216)	(2.586)	(0.962)
Atlantic	2.375*	1.562	3.719*	1.812	2.868*
	(1.298)	(1.747)	(1.942)	(2.612)	(1.472)
BC	0.192	0.055	0.456	1.074	-0.044
	(1.369)	(1.872)	(2.012)	(3.478)	(1.462)
Ontario	1.767	0.428	3.369*	0.132	1.935
	(1.209)	(1.663)	(1.739)	(2.354)	(1.362)
Prairies	1.747	-0.342	4.250**	-0.362	2.468*
	(1.314)	(1.775)	(1.931)	(2.591)	(1.497)
Quebec	0.170	-1.390	1.876	0.923	0.216
	(1.434)	(2.134)	(1.921)	(4.168)	(1.508)
Intercept	42.303***	49.222***	42.297***	41.488***	42.523***
	(1.638)	(2.256)	(2.176)	(4.141)	(1.794)
N	15973	8419	7554	2678	13295

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.4: Impacts of COVID-19 on Hours of Work of Specialists by Sex, and Practice Setting

Dep. Var: Hours of Work	All Specialists	Male	Female	Hospital	Non-Hospital
Wave1	-9.359***	-11.468***	-6.118**	3.430	-12.628***
	(2.054)	(2.803)	(2.707)	(4.172)	(2.142)
Trough1	1.327	3.011	-0.707	2.730	1.064
	(2.054)	(2.560)	(3.159)	(3.281)	(2.508)
Wave2	-1.254	-1.844	0.211	-6.146*	0.410
	(2.251)	(3.111)	(3.075)	(3.564)	(2.619)
Wave3	-0.439	-2.428	1.745	-2.492	0.307
	(1.705)	(2.072)	(2.566)	(2.950)	(1.961)
Trough2	2.836	1.866	5.229	9.783**	1.103
	(2.738)	(3.481)	(4.443)	(4.696)	(3.215)
Wave4	1.026	-1.006	4.049	4.067	0.396
	(1.639)	(2.074)	(2.530)	(2.807)	(1.939)
Wave5	-3.444	-7.925***	3.248	-6.502*	-2.208
	(2.145)	(2.725)	(3.309)	(3.384)	(2.521)
Wave5+	0.968	-0.389	2.472	4.566**	-0.038
	(1.255)	(1.632)	(1.974)	(2.203)	(1.489)
Age	-0.291***	-0.291***	-0.271***	-0.257***	-0.262***
	(0.045)	(0.055)	(0.075)	(0.069)	(0.061)
Age # Age	-0.016***	-0.015***	-0.018***	-0.004	-0.019***
	(0.003)	(0.004)	(0.005)	(0.007)	(0.004)
Male	6.048***			9.325***	4.942***
	(0.911)			(1.448)	(1.085)
Hospital	0.497	2.319*	-1.887		
	(0.951)	(1.291)	(1.350)		
Child<=5	-4.778***	-1.702	-8.624***	-6.649***	-3.978**
	(1.325)	(1.637)	(2.025)	(2.036)	(1.628)
Child>5	-0.191	0.218	-1.391	-1.189	0.095
	(1.167)	(1.448)	(1.895)	(1.929)	(1.359)
Single/Div/Sep/Widow	-0.503	-0.618	-1.101	-0.957	-0.486
	(1.117)	(1.411)	(1.703)	(1.502)	(1.416)
Immig	0.612	1.512	-0.859	-1.451	1.290
	(0.933)	(1.178)	(1.518)	(1.560)	(1.117)
Feb	-0.810	-1.124	-0.605	2.537	-1.989
	(1.122)	(1.476)	(1.739)	(2.275)	(1.318)
Mar	-1.828	-1.234	-2.347	-0.505	-2.339
	(1.314)	(1.713)	(2.089)	(2.604)	(1.513)
Apr	-0.177	-1.142	1.326	-2.049	0.422
	(1.282)	(1.661)	(2.069)	(2.671)	(1.473)
May	-1.057	-3.185*	2.064	-4.771	0.179
	(1.313)	(1.666)	(2.125)	(2.900)	(1.456)
Jun	0.006	-1.910	2.616	-1.987	0.872
	(1.462)	(1.839)	(2.363)	(2.905)	(1.683)

Jul	-3.855**	-6.183***	-0.530	-5.443*	-3.133*
	(1.545)	(1.891)	(2.543)	(2.966)	(1.792)
Aug	-7.585***	-9.910***	-4.475*	-5.757**	-7.991***
	(1.526)	(1.938)	(2.433)	(2.644)	(1.832)
Sept	-0.126	-2.119	2.618	-2.350	0.634
	(1.356)	(1.711)	(2.183)	(2.667)	(1.571)
Oct	-6.387***	-7.978***	-3.856*	-4.873*	-6.777***
	(1.314)	(1.664)	(2.089)	(2.602)	(1.527)
Nov	-1.786	-3.157*	0.273	0.631	-2.644*
	(1.276)	(1.648)	(2.048)	(2.546)	(1.453)
Dec	-0.450	-1.337	1.148	1.852	-1.117
	(1.186)	(1.470)	(1.995)	(2.465)	(1.344)
Atlantic	1.925	5.482**	-3.175	-2.755	4.145*
	(1.680)	(2.143)	(2.513)	(2.355)	(2.174)
BC	-0.849	1.431	-4.064	-5.707**	0.519
	(1.756)	(2.100)	(2.790)	(2.887)	(2.134)
Ontario	2.302	3.938**	-0.033	-0.581	2.971
	(1.522)	(1.915)	(2.355)	(2.139)	(1.931)
Prairies	-0.098	1.009	-1.643	-4.185*	1.532
	(1.618)	(2.023)	(2.568)	(2.257)	(2.093)
Quebec	1.063	1.779	0.540	-1.750	2.049
	(1.707)	(2.069)	(2.808)	(2.607)	(2.044)
Intercept	43.287***	48.070***	45.436***	43.949***	43.164***
	(1.996)	(2.325)	(3.119)	(3.009)	(2.436)
N	11037	6349	4688	2978	8059

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.5: Impacts of COVID-19 on Hours of Work by region

Dep. Var: Hours of Work	AB	Atlantic	BC	ON	MN & SK	QC
Wave1	-14.519***	-7.554***	-7.459***	-8.875***	-1.597	-1.371
	(2.582)	(2.745)	(2.770)	(2.348)	(2.724)	(2.912)
Wave1+	0.064	0.553	-0.178	0.131	1.211	-0.286
	(1.843)	(1.247)	(1.407)	(1.003)	(1.227)	(1.566)
Age	-0.130	-0.261***	-0.293***	-0.207***	-0.264***	-0.203***
	(0.085)	(0.067)	(0.074)	(0.043)	(0.055)	(0.076)
Age # Age	-0.019***	-0.008*	-0.010*	-0.015***	-0.016***	-0.017**
	(0.007)	(0.005)	(0.006)	(0.003)	(0.004)	(0.007)
Male	6.933***	9.335***	8.248***	6.342***	5.124***	4.889***
	(1.673)	(1.333)	(1.497)	(0.970)	(1.240)	(1.299)
Hospital	2.927	-1.225	0.724	1.100	-2.044*	2.786
	(1.821)	(1.289)	(2.357)	(1.066)	(1.228)	(2.136)
Child<=5	-6.279***	-2.315	-5.714***	-5.357***	-4.679**	-4.574**
	(2.293)	(1.971)	(2.205)	(1.525)	(1.936)	(1.982)
Child>5	-3.852*	3.376**	-1.312	-2.144*	-1.708	-0.453
	(2.205)	(1.436)	(1.890)	(1.299)	(1.637)	(1.674)
Single/Div/Sep/Widow	-0.211	2.592*	-0.628	-0.290	-0.229	-0.857
	(2.038)	(1.528)	(1.951)	(1.227)	(1.666)	(1.356)
Immig	1.812	0.032	-0.699	0.393	-0.403	-1.265
	(1.774)	(1.427)	(1.448)	(0.944)	(1.249)	(1.659)
Feb	-2.598	-1.621	-2.741	-1.706	-2.794*	2.032
	(2.470)	(1.297)	(1.778)	(1.101)	(1.540)	(1.575)
Mar	0.020	-2.365	-2.195	-4.179***	-0.243	1.023
	(2.121)	(1.736)	(1.952)	(1.337)	(1.624)	(1.760)
Apr	-2.966	-2.104	-2.845	-1.255	-0.609	0.503
	(2.446)	(1.548)	(1.779)	(1.235)	(1.692)	(1.661)
May	-1.640	-2.964*	-3.156*	0.223	-0.004	-2.049
	(2.469)	(1.569)	(1.859)	(1.297)	(1.789)	(1.938)
Jun	1.774	-0.857	-4.325**	-0.801	3.299*	1.782
	(2.876)	(1.670)	(1.954)	(1.376)	(1.772)	(2.011)
Jul	-6.509**	-3.640**	-5.858***	-2.824**	-2.003	-4.894**
	(3.128)	(1.820)	(2.111)	(1.435)	(1.898)	(2.089)
Aug	-10.467***	-7.182***	-7.731***	-3.558**	-4.244**	-7.283***
	(3.238)	(1.848)	(2.371)	(1.388)	(1.910)	(2.081)
Sept	-0.656	-2.416	-2.091	-0.991	0.514	1.274
	(2.811)	(1.600)	(2.054)	(1.292)	(1.713)	(2.061)
Oct	-8.530***	-4.099***	-6.589***	-5.903***	-0.681	-1.773

	(2.487)	(1.561)	(2.010)	(1.194)	(1.733)	(1.923)
Nov	-4.942**	-1.930	-3.018	-1.233	-1.706	-0.059
	(2.478)	(1.764)	(1.852)	(1.164)	(1.585)	(1.810)
Dec	-0.947	2.631	0.298	0.184	-1.613	3.025*
	(2.444)	(1.658)	(1.684)	(1.034)	(1.593)	(1.622)
Intercept	44.554***	40.013***	42.415***	44.636***	43.646***	42.543***
	(3.188)	(2.035)	(2.136)	(1.539)	(2.159)	(2.383)
N	2276	4378	3144	8622	3933	4657

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.6: Impacts of COVID-19 on Hours of Work by Immigration, and Age

Dep. Var: Hours of Work	Non-immig	Immig	28-39	40-49	50-59	60 & Above
Wave1	-7.207***	-8.399***	-5.592**	-7.407***	-7.460***	-11.105***
	(1.626)	(1.889)	(2.286)	(2.279)	(2.641)	(2.308)
Wave1+	0.022	0.164	-0.744	0.801	0.941	-0.897
	(0.764)	(1.019)	(1.016)	(1.138)	(1.153)	(1.738)
Age	-0.258***	-0.111**				
	(0.036)	(0.045)				
Age # Age	-0.013***	-0.020***				
	(0.003)	(0.003)				
Male	6.265***	7.066***	7.259***	5.981***	5.326***	3.986**
	(0.718)	(0.945)	(1.005)	(1.030)	(1.118)	(1.655)
Hospital	1.975**	0.054	3.185***	-1.259	-2.897**	2.772
	(0.919)	(1.024)	(1.082)	(1.095)	(1.296)	(2.562)
Child<=5	-5.772***	-4.016***	-5.796***	-3.925**	-1.543	-32.126***
	(1.050)	(1.512)	(1.165)	(1.756)	(4.505)	(4.125)
Child>5	-0.827	-2.527**	-1.497	-2.922**	0.768	3.398*
	(0.967)	(1.145)	(1.577)	(1.445)	(1.167)	(1.991)
Single/Div/Sep/Widow	-0.422	0.098	-0.240	-3.468***	-0.531	0.700
	(0.852)	(1.160)	(1.187)	(1.335)	(1.353)	(1.676)
Feb	-1.823**	-0.201	-2.651**	0.675	-0.651	-1.718
	(0.860)	(1.176)	(1.290)	(1.229)	(1.447)	(1.572)
Mar	-3.296***	0.058	-0.564	-4.120***	-5.261***	2.008
	(0.971)	(1.299)	(1.369)	(1.554)	(1.782)	(1.504)
Apr	-1.788**	-0.588	0.077	-1.276	-2.702*	-2.026
	(0.887)	(1.323)	(1.344)	(1.459)	(1.571)	(1.530)
May	-0.814	-1.782	-2.314	0.707	-0.436	-2.207
	(0.949)	(1.384)	(1.443)	(1.472)	(1.751)	(1.609)
Jun	-0.342	-0.035	-1.524	2.738*	-0.227	-1.595
	(1.055)	(1.377)	(1.543)	(1.509)	(1.820)	(1.797)
Jul	-4.654***	-2.643*	-3.670**	-3.532**	-4.196**	-5.170**
	(1.109)	(1.459)	(1.591)	(1.644)	(1.787)	(2.119)
Aug	-7.602***	-2.397	-4.364***	-6.001***	-5.991***	-8.076***
	(1.118)	(1.462)	(1.621)	(1.694)	(1.905)	(1.968)
Sept	-0.572	-0.601	0.086	0.243	0.368	-4.042**
	(1.022)	(1.381)	(1.550)	(1.360)	(1.709)	(2.034)
Oct	-3.683***	-7.198***	-4.746***	-3.118**	-5.611***	-6.871***
	(0.952)	(1.265)	(1.397)	(1.398)	(1.634)	(1.815)
Nov	-1.298	-2.211*	-1.653	-0.686	-1.827	-2.656
	(0.906)	(1.275)	(1.357)	(1.292)	(1.557)	(1.794)
Dec	1.454*	-0.702	1.798	1.166	1.360	-2.367
	(0.816)	(1.190)	(1.211)	(1.292)	(1.378)	(1.551)
Atlantic	3.159**	0.487	3.122*	0.734	4.499**	0.252
	(1.307)	(1.791)	(1.827)	(1.770)	(1.988)	(3.134)

BC	1.062	-2.287	2.784	-1.625	-2.156	-2.630
	(1.431)	(1.681)	(1.943)	(1.841)	(1.927)	(3.196)
Ontario	2.877**	0.978	3.300**	0.058	4.419**	-2.198
	(1.256)	(1.509)	(1.550)	(1.719)	(1.781)	(2.972)
Prairies	2.134	-0.864	1.536	1.190	2.457	-3.384
	(1.343)	(1.587)	(1.667)	(1.784)	(1.919)	(3.299)
Quebec	1.935	-1.839	1.997	0.152	3.669*	-4.105
	(1.361)	(1.925)	(1.832)	(1.942)	(1.953)	(3.415)
Immig			-2.069*	0.237	1.484	0.505
			(1.126)	(1.094)	(1.111)	(1.494)
Intercept	41.099***	43.997***	41.064***	43.983***	38.303***	38.387***
	(1.628)	(1.994)	(1.868)	(2.322)	(2.444)	(3.580)
N	18534	8476	8158	7128	6330	5394

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.7: Impacts of COVID-19 on Hours of Work by Child Status

Dep. Var: Hours of Work	No Children	Children	Children<=5 yr	Children 5+ yr
Wave1	-7.773***	-7.923***	-6.209**	-8.611***
	(1.714)	(1.779)	(2.703)	(2.243)
Wave1+	-0.128	0.172	-0.400	0.517
	(0.936)	(0.765)	(1.285)	(0.949)
Age	-0.232***	-0.007	-0.516**	-0.036
	(0.034)	(0.048)	(0.230)	(0.067)
Age # Age	-0.012***	-0.024***	-0.048***	-0.021***
	(0.003)	(0.005)	(0.013)	(0.007)
Male	5.176***	7.337***	9.552***	6.326***
	(0.879)	(0.732)	(1.261)	(0.916)
Hospital	2.788***	-0.788	-0.163	-1.425
	(1.045)	(0.921)	(1.508)	(1.117)
Single/Div/Sep/Widow	-0.253	-1.187	-3.639**	-0.068
	(0.894)	(1.121)	(1.803)	(1.383)
Immig	0.088	-0.020	0.040	-0.059
	(0.909)	(0.777)	(1.433)	(0.909)
Feb	-1.709*	-0.729	0.303	-1.496
	(1.022)	(0.953)	(1.611)	(1.180)
Mar	2.438**	-6.034***	-2.512	-8.310***
	(1.092)	(1.106)	(1.810)	(1.395)
Apr	-0.568	-1.972**	2.101	-4.639***
	(1.113)	(0.996)	(1.610)	(1.248)
May	-0.002	-1.983*	-0.561	-2.832**
	(1.216)	(1.058)	(1.688)	(1.348)
Jun	-0.614	0.184	2.863	-1.513
	(1.275)	(1.105)	(1.756)	(1.404)
Jul	-3.629***	-4.378***	-2.263	-5.622***
	(1.338)	(1.160)	(1.817)	(1.496)
Aug	-4.250***	-7.396***	-4.850***	-8.924***
	(1.333)	(1.181)	(1.832)	(1.535)
Sept	-1.811	0.399	0.826	0.116
	(1.236)	(1.083)	(1.756)	(1.376)
Oct	-5.689***	-4.312***	-3.718**	-4.601***
	(1.151)	(1.007)	(1.639)	(1.252)
Nov	-0.943	-2.332**	-1.254	-2.996**
	(1.120)	(0.953)	(1.423)	(1.263)
Dec	1.153	0.258	0.912	-0.328
	(1.011)	(0.887)	(1.367)	(1.161)
Atlantic	0.247	3.851***	3.285	4.060**
	(1.603)	(1.345)	(2.291)	(1.636)
BC	-0.453	-0.385	0.365	-0.832
	(1.727)	(1.350)	(2.281)	(1.636)

Ontario	1.390	2.209*	3.059	1.630
	(1.484)	(1.227)	(2.038)	(1.498)
Prairies	-0.331	2.089	2.805	1.610
	(1.565)	(1.327)	(2.257)	(1.611)
Quebec	-0.509	2.100	3.652	1.533
	(1.762)	(1.349)	(2.227)	(1.635)
Intercept	42.372***	40.784***	35.192***	42.701***
	(1.759)	(1.500)	(2.815)	(1.868)
N	12744	14266	5478	8788

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

TABLE A2.8: Impacts of COVID-19 on Hours of Work by Sex & Child Status

Dep. Var: Hours of Work	Male W/O Child	Female W/O Child	Male W/ Child	Female W/ Child
Wave1	-7.804***	-7.970***	-8.121***	-7.413***
	(2.484)	(2.218)	(2.648)	(2.207)
Wave1+	0.473	-0.714	-0.360	0.538
	(1.304)	(1.306)	(1.038)	(1.107)
Age	-0.157***	-0.329***	-0.156***	0.180**
	(0.050)	(0.046)	(0.056)	(0.085)
Age # Age	-0.013***	-0.014***	-0.018***	-0.025***
	(0.003)	(0.003)	(0.005)	(0.008)
Hospital	4.730***	0.910	-0.963	-0.945
	(1.629)	(1.207)	(1.310)	(1.229)
Single/Div/Sep/Widow	0.154	-0.191	-3.778**	-0.442
	(1.308)	(1.194)	(1.466)	(1.512)
Immig	0.342	-0.805	0.220	-0.234
	(1.235)	(1.326)	(1.047)	(1.140)
Feb	-1.117	-2.355	-0.153	-1.444
	(1.371)	(1.510)	(1.368)	(1.321)
Mar	2.640*	2.610	-5.699***	-6.573***
	(1.464)	(1.650)	(1.568)	(1.561)
Apr	-0.348	-0.675	-2.076	-2.120
	(1.541)	(1.553)	(1.328)	(1.467)
May	-0.837	1.480	-1.483	-2.585*
	(1.691)	(1.662)	(1.423)	(1.538)
Jun	-1.492	0.762	-0.536	0.714
	(1.779)	(1.782)	(1.486)	(1.607)
Jul	-5.043***	-1.692	-3.684**	-5.111***
	(1.813)	(1.927)	(1.584)	(1.676)
Aug	-5.821***	-2.249	-7.310***	-7.622***
	(1.741)	(2.035)	(1.630)	(1.697)
Sept	-3.486**	0.307	0.008	0.590
	(1.735)	(1.677)	(1.464)	(1.573)
Oct	-6.931***	-4.100**	-4.892***	-3.918***
	(1.562)	(1.661)	(1.377)	(1.446)
Nov	-1.763	0.275	-1.840	-2.958**
	(1.545)	(1.611)	(1.376)	(1.307)
Dec	-0.097	3.005**	1.334	-0.994
	(1.352)	(1.515)	(1.226)	(1.269)
Atlantic	-0.700	0.483	6.471***	1.348
	(2.176)	(2.295)	(1.750)	(2.036)
BC	-1.567	0.587	1.814	-3.000
	(2.310)	(2.484)	(1.662)	(2.154)
Ontario	-0.220	3.077	3.185**	1.248
	(2.065)	(2.038)	(1.577)	(1.891)

Prairies	-2.229	1.950	1.862	2.589
	(2.117)	(2.242)	(1.706)	(2.044)
Quebec	-3.095	2.320	3.112*	1.454
	(2.402)	(2.453)	(1.720)	(2.029)
Intercept	48.927***	40.801***	46.575***	42.717***
	(2.402)	(2.311)	(1.709)	(2.233)
N	7433	5311	7335	6931

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.9: Impacts of COVID-19 on Hours of Work by Marital & Child Status

Dep. Var: Hours of Work	Married/CL W/O Child	Sing/Div/Sep/Widow W/O Child	Married/CL W/ Child	Other W/ Child
Wave1	-8.355***	-6.995**	-9.547***	1.209
	(2.156)	(2.754)	(1.967)	(2.748)
Wave1+	1.105	-1.389	0.454	-1.370
	(1.301)	(1.339)	(0.831)	(1.922)
Age	-0.241***	-0.238***	-0.031	0.088
	(0.047)	(0.049)	(0.053)	(0.098)
Age # Age	-0.012***	-0.013***	-0.027***	-0.011
	(0.003)	(0.004)	(0.005)	(0.009)
Male	5.718***	4.406***	8.062***	3.741**
	(1.191)	(1.279)	(0.810)	(1.694)
Hospital	1.814	3.427**	-0.854	-0.184
	(1.465)	(1.478)	(1.012)	(2.108)
Immig	1.066	-1.401	-0.436	3.100
	(1.167)	(1.454)	(0.821)	(2.346)
Feb	-1.811	-1.645	-0.706	-1.171
	(1.370)	(1.554)	(1.033)	(2.406)
Mar	2.072	2.710	-5.874***	-6.615**
	(1.424)	(1.703)	(1.196)	(2.837)
Apr	-1.936	1.083	-1.465	-4.293
	(1.368)	(1.838)	(1.055)	(2.762)
May	-0.119	-0.073	-1.212	-5.993**
	(1.587)	(1.863)	(1.127)	(2.921)
Jun	-1.082	-0.286	0.390	-1.275
	(1.621)	(2.026)	(1.157)	(3.300)
Jul	-4.340***	-2.894	-4.595***	-3.661
	(1.601)	(2.219)	(1.251)	(3.068)
Aug	-5.613***	-2.634	-7.786***	-5.960**
	(1.688)	(2.120)	(1.292)	(2.966)
Sept	-2.118	-1.439	0.359	0.148
	(1.597)	(1.900)	(1.152)	(2.981)
Oct	-7.425***	-3.671**	-4.500***	-3.347
	(1.463)	(1.800)	(1.058)	(3.004)
Nov	-2.947**	1.401	-2.687***	-0.274
	(1.482)	(1.679)	(1.009)	(2.712)
Dec	0.320	2.202	-0.094	2.339
	(1.270)	(1.612)	(0.938)	(2.562)
Atlantic	-2.353	3.891	4.046***	1.757
	(1.991)	(2.640)	(1.437)	(3.365)
BC	-2.169	1.864	0.437	-9.493***
	(2.096)	(2.874)	(1.431)	(3.454)
Ontario	0.289	2.912	2.528*	-0.185
	(1.865)	(2.389)	(1.311)	(3.253)

Prairies	-1.844	1.487	2.306	0.274
	(2.019)	(2.426)	(1.402)	(4.127)
Quebec	-3.239	2.517	3.577**	-2.514
	(2.418)	(2.614)	(1.533)	(2.388)
Intercept	43.734***	40.391***	40.203***	43.560***
	(2.336)	(2.699)	(1.632)	(3.178)
N	7206	5538	11950	2316

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.10: Logit estimates of COVID-19 and Absences of All Physicians by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	All	Male	Female	Hospital	Non-Hospital
Wave1	0.027	0.044*	0.009	-0.047**	0.044**
	(0.016)	(0.023)	(0.023)	(0.021)	(0.019)
Trough1	-0.012	-0.029**	0.013	-0.001	-0.016
	(0.013)	(0.013)	(0.022)	(0.028)	(0.013)
Wave2	-0.020	-0.030	-0.012	0.025	-0.029
	(0.017)	(0.023)	(0.024)	(0.044)	(0.018)
Wave3	-0.041***	-0.043***	-0.039**	-0.006	-0.047***
	(0.011)	(0.014)	(0.018)	(0.032)	(0.012)
Trough2	0.013	-0.004	0.035	-0.001	0.014
	(0.018)	(0.018)	(0.034)	(0.039)	(0.019)
Wave4	-0.020	-0.014	-0.025	-0.007	-0.024
	(0.013)	(0.017)	(0.020)	(0.030)	(0.014)
Wave5	-0.024*	-0.015	-0.034*	-0.023	-0.024
	(0.013)	(0.019)	(0.019)	(0.024)	(0.016)
Wave5+	-0.001	-0.013	0.011	0.020	-0.007
	(0.010)	(0.011)	(0.016)	(0.022)	(0.010)
Age	0.001*	0.001*	-0.000	0.001	0.000
	(0.000)	(0.000)	(0.001)	(0.001)	(0.000)
Male	-0.030***			-0.052***	-0.024***
	(0.006)			(0.014)	(0.007)
Hospital	0.003	-0.006	0.015		
	(0.008)	(0.010)	(0.014)		
Child<=5	0.062***	0.004	0.117***	0.117***	0.044***
	(0.012)	(0.013)	(0.019)	(0.027)	(0.013)
Child>5	0.005	-0.002	0.016	0.031*	-0.002
	(0.008)	(0.011)	(0.011)	(0.017)	(0.009)
Single/Div/Sep/Widow	0.009	0.000	0.022**	0.009	0.010
	(0.007)	(0.010)	(0.011)	(0.016)	(0.008)
Immig	-0.006	-0.009	0.001	0.010	-0.010
	(0.007)	(0.009)	(0.011)	(0.015)	(0.008)
Feb	0.031***	0.020	0.045***	0.022	0.034***
	(0.010)	(0.013)	(0.017)	(0.022)	(0.012)
Mar	0.070***	0.054***	0.088***	0.056**	0.073***
	(0.012)	(0.015)	(0.020)	(0.027)	(0.014)
Apr	-0.002	-0.019*	0.018	-0.020	0.003
	(0.009)	(0.011)	(0.015)	(0.023)	(0.010)
May	0.017*	0.006	0.027*	0.024	0.016
	(0.010)	(0.013)	(0.016)	(0.026)	(0.011)
Jun	0.015	0.008	0.022	0.027	0.012

	(0.011)	(0.014)	(0.016)	(0.027)	(0.011)
Jul	0.071***	0.067***	0.076***	0.086***	0.068***
	(0.012)	(0.016)	(0.018)	(0.030)	(0.013)
Aug	0.108***	0.105***	0.113***	0.055**	0.122***
	(0.013)	(0.018)	(0.020)	(0.027)	(0.015)
Sept	0.018*	0.026*	0.008	-0.004	0.023**
	(0.010)	(0.014)	(0.015)	(0.025)	(0.011)
Oct	0.037***	0.035**	0.039**	0.043	0.035***
	(0.011)	(0.014)	(0.017)	(0.028)	(0.012)
Nov	0.025**	0.019	0.032*	-0.002	0.031***
	(0.010)	(0.013)	(0.016)	(0.024)	(0.011)
Dec	-0.006	-0.010	-0.001	-0.017	-0.004
	(0.009)	(0.011)	(0.015)	(0.021)	(0.010)
Atlantic	-0.019	-0.028*	-0.007	-0.001	-0.024
	(0.013)	(0.016)	(0.021)	(0.023)	(0.015)
BC	-0.016	-0.026	0.001	0.033	-0.026*
	(0.014)	(0.016)	(0.023)	(0.033)	(0.015)
Ontario	-0.022*	-0.024	-0.018	0.008	-0.028**
	(0.012)	(0.015)	(0.020)	(0.023)	(0.014)
Prairies	-0.019	-0.013	-0.025	0.005	-0.026*
	(0.013)	(0.016)	(0.021)	(0.024)	(0.015)
Quebec	-0.004	-0.001	-0.006	0.018	-0.011
	(0.014)	(0.017)	(0.021)	(0.031)	(0.015)
N	27010	14768	12242	5656	21354

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.11: Logit estimates of COVID-19 and Absences of GP/FPs by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	All	Male	Female	Hospital	Non-Hospital
Wave1	0.028	0.048	0.013	-0.070***	0.046*
	(0.021)	(0.034)	(0.026)	(0.024)	(0.025)
Trough1	-0.025	-0.042**	0.000	-0.051	-0.021
	(0.016)	(0.017)	(0.027)	(0.036)	(0.017)
Wave2	-0.041**	-0.075***	-0.004	-0.053	-0.040**
	(0.018)	(0.014)	(0.032)	(0.048)	(0.019)
Wave3	-0.042***	-0.054***	-0.030	-0.014	-0.045***
	(0.015)	(0.019)	(0.022)	(0.050)	(0.015)
Trough2	0.024	-0.008	0.065	0.056	0.021
	(0.025)	(0.025)	(0.044)	(0.072)	(0.025)
Wave4	-0.005	-0.021	0.021	0.018	-0.010
	(0.018)	(0.022)	(0.029)	(0.047)	(0.020)
Wave5	-0.035**	-0.053**	-0.016	-0.065**	-0.028
	(0.016)	(0.021)	(0.024)	(0.026)	(0.019)
Wave5+	-0.010	-0.039***	0.015	-0.004	-0.013
	(0.012)	(0.013)	(0.020)	(0.035)	(0.012)
Age	0.000	0.001	-0.001	-0.001	-0.000
	(0.000)	(0.001)	(0.001)	(0.002)	(0.001)
Male	-0.025***			-0.025	-0.024***
	(0.008)			(0.022)	(0.009)
Hospital	0.013	0.019	0.011		
	(0.013)	(0.017)	(0.019)		
Child<=5	0.056***	0.009	0.099***	0.114***	0.041**
	(0.016)	(0.018)	(0.024)	(0.040)	(0.016)
Child>5	0.014	0.012	0.017	0.052*	0.008
	(0.010)	(0.015)	(0.014)	(0.029)	(0.011)
Single/Div/Sep/Widow	-0.003	-0.013	0.010	-0.008	-0.000
	(0.009)	(0.012)	(0.013)	(0.024)	(0.010)
Immig	-0.011	-0.009	-0.010	0.002	-0.013
	(0.009)	(0.012)	(0.014)	(0.022)	(0.010)
Feb	0.018	0.004	0.033	0.045	0.013
	(0.014)	(0.017)	(0.021)	(0.039)	(0.014)
Mar	0.061***	0.038*	0.085***	0.071*	0.057***
	(0.016)	(0.020)	(0.025)	(0.039)	(0.017)
Apr	-0.001	-0.026*	0.028	-0.029	0.004
	(0.012)	(0.015)	(0.019)	(0.031)	(0.014)
May	0.025*	-0.000	0.050**	0.023	0.025
	(0.014)	(0.018)	(0.021)	(0.035)	(0.016)
Jun	0.011	-0.010	0.032	0.034	0.006

	(0.014)	(0.018)	(0.020)	(0.039)	(0.015)
Jul	0.076***	0.059***	0.094***	0.118***	0.069***
	(0.016)	(0.022)	(0.023)	(0.044)	(0.017)
Aug	0.086***	0.079***	0.094***	0.053	0.095***
	(0.017)	(0.024)	(0.025)	(0.039)	(0.020)
Sept	0.021	0.043**	0.002	-0.009	0.027*
	(0.014)	(0.021)	(0.019)	(0.037)	(0.016)
Oct	0.028*	0.024	0.033	0.075	0.017
	(0.015)	(0.020)	(0.021)	(0.046)	(0.015)
Nov	0.024*	0.016	0.032	0.028	0.023
	(0.014)	(0.019)	(0.021)	(0.041)	(0.015)
Dec	-0.027**	-0.026*	-0.028*	-0.012	-0.030***
	(0.011)	(0.014)	(0.016)	(0.032)	(0.011)
Atlantic	-0.021	-0.031*	-0.011	-0.029	-0.020
	(0.016)	(0.018)	(0.027)	(0.039)	(0.017)
BC	-0.007	-0.015	0.004	0.030	-0.012
	(0.017)	(0.019)	(0.030)	(0.050)	(0.018)
Ontario	-0.016	-0.011	-0.022	-0.021	-0.015
	(0.016)	(0.018)	(0.026)	(0.037)	(0.017)
Prairies	-0.017	-0.004	-0.029	-0.008	-0.020
	(0.017)	(0.020)	(0.026)	(0.039)	(0.018)
Quebec	0.016	0.032	0.004	0.041	0.010
	(0.018)	(0.023)	(0.029)	(0.055)	(0.019)
N	15973	8419	7554	2678	13295

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.12: Logit estimates of COVID-19 and Absences of Specialists by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	All	Male	Female	Hospital	Non-Hospital
Wave1	0.023	0.038	-0.001	-0.031	0.042
	(0.025)	(0.028)	(0.048)	(0.034)	(0.031)
Trough1	0.006	-0.016	0.036	0.026	-0.006
	(0.020)	(0.019)	(0.038)	(0.034)	(0.022)
Wave2	0.006	0.028	-0.027	0.075	-0.009
	(0.029)	(0.044)	(0.035)	(0.055)	(0.034)
Wave3	-0.041**	-0.030	-0.055*	0.009	-0.053***
	(0.017)	(0.020)	(0.029)	(0.046)	(0.018)
Trough2	-0.007	-0.002	-0.026	-0.052**	-0.001
	(0.024)	(0.027)	(0.040)	(0.022)	(0.028)
Wave4	-0.048***	-0.014	-0.094***	-0.052**	-0.048**
	(0.018)	(0.027)	(0.022)	(0.025)	(0.021)
Wave5	-0.006	0.036	-0.063*	0.026	-0.018
	(0.025)	(0.035)	(0.032)	(0.042)	(0.027)
Wave5+	0.010	0.020	-0.001	0.033	0.002
	(0.016)	(0.019)	(0.025)	(0.026)	(0.018)
Age	0.001***	0.001**	0.000	0.002**	0.001
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)
Male	-0.037***			-0.078***	-0.026**
	(0.010)			(0.019)	(0.011)
Hospital	-0.010	-0.025**	0.014		
	(0.011)	(0.011)	(0.020)		
Child<=5	0.069***	0.005	0.144***	0.127***	0.046**
	(0.018)	(0.018)	(0.030)	(0.036)	(0.019)
Child>5	-0.006	-0.019	0.017	0.017	-0.016
	(0.012)	(0.016)	(0.018)	(0.018)	(0.014)
Single/Div/Sep/Widow	0.028**	0.021	0.044**	0.019	0.029**
	(0.012)	(0.015)	(0.018)	(0.019)	(0.014)
Immig	-0.001	-0.013	0.016	0.019	-0.007
	(0.011)	(0.011)	(0.019)	(0.019)	(0.012)
Feb	0.050***	0.039**	0.069***	0.004	0.068***
	(0.016)	(0.020)	(0.027)	(0.024)	(0.021)
Mar	0.084***	0.074***	0.096***	0.041	0.101***
	(0.019)	(0.023)	(0.033)	(0.035)	(0.023)
Apr	-0.004	-0.011	0.005	-0.008	0.001
	(0.014)	(0.015)	(0.025)	(0.032)	(0.015)
May	0.005	0.014	-0.007	0.033	0.001
	(0.015)	(0.018)	(0.025)	(0.036)	(0.016)
Jun	0.022	0.033*	0.011	0.025	0.022

	(0.016)	(0.020)	(0.027)	(0.037)	(0.018)
Jul	0.064***	0.078***	0.050*	0.064	0.067***
	(0.018)	(0.023)	(0.028)	(0.039)	(0.020)
Aug	0.136***	0.141***	0.137***	0.060	0.164***
	(0.020)	(0.025)	(0.033)	(0.038)	(0.024)
Sept	0.013	0.013	0.015	0.008	0.018
	(0.015)	(0.018)	(0.025)	(0.034)	(0.016)
Oct	0.050***	0.052***	0.047*	0.022	0.062***
	(0.016)	(0.020)	(0.027)	(0.034)	(0.018)
Nov	0.027*	0.028	0.027	-0.023	0.046***
	(0.015)	(0.017)	(0.026)	(0.028)	(0.017)
Dec	0.025	0.013	0.044	-0.023	0.042**
	(0.016)	(0.017)	(0.028)	(0.026)	(0.019)
Atlantic	-0.016	-0.036	0.006	0.029	-0.033
	(0.022)	(0.028)	(0.034)	(0.026)	(0.029)
BC	-0.031	-0.053*	-0.002	0.031	-0.053**
	(0.022)	(0.028)	(0.033)	(0.036)	(0.026)
Ontario	-0.033	-0.050*	-0.010	0.035	-0.054**
	(0.020)	(0.027)	(0.031)	(0.025)	(0.025)
Prairies	-0.024	-0.032	-0.015	0.021	-0.041
	(0.022)	(0.028)	(0.035)	(0.026)	(0.028)
Quebec	-0.031	-0.047	-0.018	0.015	-0.049*
	(0.022)	(0.029)	(0.033)	(0.029)	(0.026)
N	11037	6349	4688	2978	8059

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.13: NB Estimates of COVID-19 and Length of Absences by Sex, and Practice Setting

Average Marginal Effect (in week)	All	Male	Female	Hospital	Non-Hospital
Wave1	-0.105	2.079*	-3.098*	2.601	-0.652
	(1.260)	(1.215)	(1.836)	(5.645)	(1.220)
Trough1	-1.068	0.910	-3.926***	1.628	-0.720
	(1.261)	(1.511)	(1.497)	(3.562)	(1.466)
Wave2	2.926	2.110	1.388	11.128*	1.644
	(2.600)	(2.285)	(3.037)	(6.281)	(2.764)
Wave3	0.905	-0.291	3.253	12.411	-1.744
	(3.000)	(2.009)	(6.578)	(11.394)	(2.044)
Trough2	-3.890***	-1.664*	-6.025***	-3.831**	-3.612***
	(0.877)	(0.929)	(1.372)	(1.769)	(0.963)
Wave4	1.545	2.375	-1.442	-2.412	2.300
	(3.253)	(3.187)	(3.344)	(2.152)	(3.829)
Wave5	2.400	3.171	-2.095	-1.718	3.368
	(3.511)	(3.290)	(2.564)	(2.431)	(4.199)
Wave5+	-0.201	0.041	0.285	2.793	-0.847
	(1.315)	(1.229)	(2.275)	(3.771)	(1.288)
Age	0.159***	0.196***	0.120	0.123	0.172***
	(0.039)	(0.030)	(0.081)	(0.086)	(0.041)
Male	-2.737***			-4.071***	-2.479***
	(0.732)			(1.306)	(0.776)
Hospital	1.504	0.431	3.094**		
	(1.041)	(0.907)	(1.575)		
Child<=5	9.323***	0.669	14.361***	9.913***	9.319***
	(1.477)	(0.824)	(2.178)	(3.134)	(1.665)
Child>5	1.521*	1.676	1.491	-0.008	1.631**
	(0.815)	(1.026)	(0.926)	(1.404)	(0.828)
Single/Div/Sep/Widow	1.378	0.653	3.416**	0.551	1.159
	(0.894)	(0.791)	(1.386)	(1.784)	(0.907)
Immig	-0.316	-0.219	0.128	-0.898	0.064
	(0.714)	(0.594)	(1.169)	(1.482)	(0.741)
Feb	-1.683*	-2.046**	-0.608	0.077	-2.167**
	(0.862)	(1.006)	(1.456)	(1.491)	(0.971)
Mar	-1.826*	-2.290**	-0.638	0.312	-2.147**
	(0.948)	(1.061)	(1.745)	(1.709)	(1.060)
Apr	0.755	-2.483**	2.876	0.342	0.780
	(1.425)	(1.176)	(2.205)	(2.468)	(1.579)
May	1.554	-1.305	4.128**	2.358	1.430
	(1.257)	(1.201)	(2.098)	(2.388)	(1.347)
Jun	0.392	-1.530	1.683	-0.314	1.151
	(1.226)	(1.183)	(1.923)	(2.150)	(1.406)
Jul	0.037	-1.781	1.560	4.101	-0.970
	(1.115)	(1.178)	(1.780)	(2.606)	(1.157)

Aug	-0.312	-1.831*	1.464	2.823	-1.138
	(1.077)	(1.093)	(1.783)	(2.520)	(1.099)
Sept	0.968	-0.542	2.039	2.967	0.441
	(1.266)	(1.278)	(1.826)	(2.612)	(1.340)
Oct	1.100	-1.036	4.646*	1.061	1.278
	(1.243)	(1.154)	(2.404)	(2.211)	(1.386)
Nov	1.806	1.418	2.476	4.689*	1.638
	(1.525)	(1.851)	(1.901)	(2.718)	(1.686)
Dec	1.750	1.228	2.367	1.641	2.251
	(1.511)	(1.680)	(2.014)	(2.398)	(1.755)
Atlantic	1.199	-0.185	3.022	-0.843	0.255
	(1.610)	(1.215)	(2.889)	(5.541)	(1.532)
BC	-0.197	-1.183	-0.141	-5.489	-0.404
	(1.158)	(0.864)	(2.071)	(5.268)	(1.144)
Ontario	-0.205	0.004	-1.440	-5.029	-0.376
	(1.008)	(0.897)	(1.566)	(5.110)	(0.998)
Prairies	-0.286	-0.202	0.559	-4.294	0.217
	(1.127)	(1.004)	(2.101)	(5.583)	(1.164)
Quebec	0.338	0.090	0.100	-7.138	0.844
	(1.105)	(1.116)	(1.717)	(5.089)	(1.174)
Year2016	-1.475	-1.762	-0.822	0.485	-1.968
	(1.316)	(1.280)	(2.120)	(2.772)	(1.442)
Year2017	-2.920**	-2.627**	-3.560*	-3.170	-2.448*
	(1.275)	(1.187)	(1.876)	(2.433)	(1.345)
Year2018	-2.188*	-1.703	-3.465**	0.243	-2.774**
	(1.122)	(1.124)	(1.724)	(2.503)	(1.183)
N	2704	1295	1409	548	2156

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.14: NB Estimates of COVID-19 and Length of Absences of GP/FPs by Sex, and Practice Setting

Average Marginal Effect (in week)	All	Male	Female	Hospital	Non-Hospital
Wave1	-0.319	1.595	-4.264*	3.903	-0.989
	(1.453)	(1.304)	(2.312)	(7.637)	(1.430)
Trough1	-1.864	-0.725	-5.055**	7.320	-2.226
	(1.431)	(1.772)	(1.995)	(7.580)	(1.508)
Wave2	0.128	1.003	-3.057	20.383*	-1.338
	(1.950)	(2.033)	(2.555)	(10.839)	(1.796)
Wave3	-0.340	0.588	-2.082	0.916	-0.720
	(2.884)	(3.310)	(4.085)	(3.913)	(2.838)
Trough2	-4.135***	-1.654	-6.615***	-2.069	-4.117***
	(1.131)	(1.408)	(1.975)	(2.082)	(1.247)
Wave4	2.113	4.720	-2.373	-2.742	3.360
	(3.806)	(5.152)	(3.341)	(1.840)	(4.528)
Wave5	4.314	8.600	-2.150	-2.249	4.830
	(5.521)	(9.228)	(3.278)	(2.289)	(6.125)
Wave5+	0.927	-1.552	2.649	3.117	0.201
	(1.889)	(1.126)	(3.566)	(3.677)	(1.933)
Age	0.177***	0.245***	0.099	0.060	0.201***
	(0.048)	(0.046)	(0.107)	(0.118)	(0.050)
Male	-2.088**			-3.087**	-2.016**
	(0.913)			(1.452)	(1.010)
Hospital	1.181	0.649	1.455		
	(1.293)	(1.374)	(1.724)		
Child<=5	9.558***	2.147*	14.402***	7.841**	9.632***
	(1.881)	(1.292)	(3.239)	(3.339)	(2.106)
Child>5	2.143**	3.758**	1.235	0.607	2.212*
	(1.057)	(1.762)	(1.002)	(1.387)	(1.156)
Single/Div/Sep/Widow	0.810	0.601	3.901**	2.792	0.262
	(1.005)	(1.175)	(1.726)	(2.293)	(1.061)
Immig	-0.752	-0.128	-0.285	0.315	-0.660
	(0.871)	(0.849)	(1.350)	(1.630)	(0.945)
Feb	-1.125	-1.203	-1.105	-2.070	-1.373
	(1.133)	(1.342)	(1.778)	(2.996)	(1.273)
Mar	-0.791	-1.539	0.152	0.258	-1.168
	(1.190)	(1.404)	(2.341)	(3.232)	(1.312)
Apr	-0.164	-2.341	0.810	-3.956	-0.012
	(1.464)	(1.548)	(2.174)	(3.601)	(1.581)
May	1.395	-1.440	2.443	-0.677	1.325
	(1.430)	(1.388)	(2.288)	(4.483)	(1.513)
Jun	2.083	0.066	2.321	-3.463	3.570*
	(1.544)	(1.506)	(2.312)	(3.747)	(1.960)

Jul	0.067	-1.917	0.717	-1.059	-0.620
	(1.319)	(1.519)	(1.944)	(3.931)	(1.402)
Aug	0.408	-1.835	1.736	-1.122	-0.441
	(1.435)	(1.370)	(2.330)	(3.759)	(1.412)
Sept	0.257	-1.107	0.481	-2.346	0.112
	(1.428)	(1.494)	(2.062)	(3.743)	(1.551)
Oct	2.225	-0.097	4.805	-1.549	2.779
	(1.681)	(1.494)	(3.362)	(3.663)	(1.922)
Nov	0.567	-0.124	0.527	-1.606	0.491
	(1.360)	(1.480)	(2.023)	(3.549)	(1.530)
Dec	2.922	2.228	2.184	1.691	3.853
	(2.363)	(2.221)	(2.794)	(4.367)	(2.862)
Atlantic	-1.314	-0.950	-0.492	-7.182	-1.848
	(1.446)	(1.376)	(2.883)	(7.427)	(1.626)
BC	-0.978	-0.997	-2.323	-6.926	-1.893
	(1.412)	(1.255)	(2.292)	(7.454)	(1.449)
Ontario	-0.784	0.131	-3.141	-10.695	-0.868
	(1.365)	(1.271)	(2.096)	(7.138)	(1.498)
Prairies	-0.947	0.177	-0.937	-9.430	-0.748
	(1.468)	(1.436)	(2.810)	(7.435)	(1.641)
Quebec	0.726	0.904	-0.229	-11.748	0.849
	(1.587)	(1.750)	(2.262)	(7.300)	(1.737)
Year2016	-2.410	-1.412	-3.824	3.633	-3.308*
	(1.789)	(1.757)	(2.577)	(3.412)	(1.977)
Year2017	-4.723***	-5.181***	-4.535*	-3.407	-4.341**
	(1.679)	(1.526)	(2.530)	(2.556)	(1.823)
Year2018	-1.543	-2.402*	-2.537	0.809	-2.277
	(1.361)	(1.334)	(2.268)	(2.662)	(1.495)
N	1505	701	804	259	1246

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.15: NB Estimates of COVID-19 and Length of Absences of Specialists by Sex, and Practice Setting

Average Marginal Effect (in week)	All	Male	Female	Hospital	Non-Hospital
Wave1	1.317	3.995*	-1.577	-3.940	1.803
	(1.818)	(2.052)	(2.301)	(2.625)	(1.772)
Trough1	0.367	2.829*	-2.194	-2.462	2.462
	(1.767)	(1.712)	(2.450)	(2.933)	(2.435)
Wave2	9.099	4.294	13.455	15.767	7.778
	(5.691)	(2.942)	(9.574)	(10.973)	(6.063)
Wave3	6.116	-0.525	20.630	21.223	-3.207***
	(7.749)	(0.641)	(22.046)	(19.517)	(0.917)
Trough2	-2.934***	-0.743	-4.943**	-5.365**	-1.989*
	(0.989)	(0.499)	(1.957)	(2.167)	(1.081)
Wave4	-2.170	-0.395	-2.526	0.803	-2.303**
	(1.388)	(0.550)	(2.898)	(5.086)	(1.131)
Wave5	1.160	2.113*	-2.817	0.148	1.268
	(1.665)	(1.123)	(2.851)	(5.767)	(1.388)
Wave5+	-0.245	2.911*	-3.130	3.106	-0.375
	(1.499)	(1.487)	(2.066)	(5.346)	(1.236)
Age	0.099*	0.150***	-0.031	0.032	0.106**
	(0.057)	(0.033)	(0.115)	(0.133)	(0.050)
Male	-3.235***			-4.978***	-2.824***
	(0.900)			(1.803)	(0.810)
Hospital	1.727	0.556	3.340		
	(1.264)	(0.700)	(2.097)		
Child<=5	10.690***	-0.074	17.110***	6.560	9.450***
	(2.366)	(0.580)	(3.408)	(4.579)	(2.084)
Child>5	0.625	-0.582	2.129	-3.765	0.915
	(0.839)	(0.485)	(1.333)	(2.404)	(0.788)
Single/Div/Sep/Widow	2.302**	0.816	4.373**	-4.580	2.864***
	(1.151)	(0.582)	(2.082)	(2.860)	(1.014)
Immig	0.021	0.074	0.495	-4.273**	0.923
	(0.874)	(0.445)	(1.634)	(2.119)	(0.803)
Feb	-1.851	-2.789**	0.758	0.259	-3.526**
	(1.244)	(1.228)	(1.802)	(1.498)	(1.616)
Mar	-2.571**	-2.426**	-1.227	-0.231	-3.898**
	(1.244)	(1.237)	(1.674)	(2.240)	(1.645)
Apr	2.672	-2.741**	7.937**	3.973	0.411
	(2.411)	(1.368)	(3.929)	(4.077)	(2.468)
May	3.336	-1.029	8.685**	8.674**	1.704
	(2.297)	(1.669)	(4.065)	(3.715)	(2.617)
Jun	-0.797	-2.596*	1.710	2.882	-1.992
	(1.643)	(1.352)	(2.527)	(3.065)	(1.909)

Jul	0.633	-1.508	3.744	9.169**	-2.271
	(1.558)	(1.307)	(2.514)	(4.066)	(1.721)
Aug	0.274	-1.439	3.067	3.745	-1.777
	(1.457)	(1.266)	(2.292)	(2.523)	(1.729)
Sept	2.582	-0.467	5.953**	7.986**	0.063
	(1.893)	(1.385)	(2.911)	(3.424)	(2.039)
Oct	0.479	-1.625	5.982*	4.244	-1.259
	(1.457)	(1.231)	(3.255)	(3.742)	(1.641)
Nov	4.954*	1.975	9.717**	11.478**	2.439
	(3.008)	(2.499)	(4.361)	(5.795)	(3.235)
Dec	1.416	0.288	4.074	0.690	0.604
	(1.681)	(1.564)	(2.668)	(2.307)	(2.080)
Atlantic	4.889*	1.034	6.854	5.689	3.652
	(2.677)	(1.479)	(4.773)	(5.236)	(2.568)
BC	0.029	-1.106	1.830	-1.354	0.332
	(1.369)	(0.731)	(3.208)	(3.566)	(1.337)
Ontario	0.600	-0.391	0.605	1.830	-0.269
	(1.016)	(0.618)	(1.721)	(3.386)	(0.852)
Prairies	0.804	0.923	3.352	1.691	0.934
	(1.279)	(0.983)	(3.135)	(5.022)	(1.239)
Quebec	0.160	-0.700	0.351	-1.667	0.446
	(1.220)	(0.644)	(2.354)	(3.402)	(1.187)
Year2016	0.081	-0.640	2.443	-0.924	0.328
	(1.508)	(0.576)	(2.930)	(3.130)	(1.464)
Year2017	0.142	1.176*	-2.110	-0.712	1.000
	(1.329)	(0.703)	(2.390)	(4.213)	(1.229)
Year2018	-3.855***	-0.519	-7.427***	-6.492	-3.210***
	(1.372)	(0.567)	(2.458)	(3.999)	(1.236)
N	1199	594	605	289	910

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Table A2.16: Logit estimates of COVID-19 and Exiting the LF by Sex, and Practice Setting

Average Marginal Effect (Δ Probability)	All	Male	Female	Hospital	Non-Hospital
Wave1	0.016	0.019	0.009	-0.033**	0.026*
	(0.012)	(0.018)	(0.013)	(0.015)	(0.013)
Wave1+	-0.003	-0.005	-0.002	-0.023**	0.002
	(0.005)	(0.006)	(0.007)	(0.012)	(0.005)
Age	0.002***	0.002***	0.002***	0.001***	0.002***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Male	-0.005			-0.002	-0.005
	(0.005)			(0.010)	(0.005)
Hospital	0.028***	0.034***	0.022*		
	(0.009)	(0.012)	(0.012)		
Child<=5	0.001	-0.015	0.016	0.001	0.000
	(0.009)	(0.011)	(0.012)	(0.016)	(0.010)
Child>5	-0.017***	-0.019***	-0.013**	-0.024**	-0.015***
	(0.004)	(0.007)	(0.005)	(0.012)	(0.005)
Single/Div/Sep/Widow	-0.011***	-0.015***	-0.007	-0.038***	-0.004
	(0.004)	(0.006)	(0.006)	(0.009)	(0.005)
Immig	-0.001	-0.000	0.001	0.008	-0.003
	(0.004)	(0.007)	(0.006)	(0.010)	(0.004)
Feb	-0.003	-0.005	-0.000	-0.022**	0.003
	(0.004)	(0.006)	(0.005)	(0.010)	(0.004)
Mar	-0.007	-0.008	-0.005	-0.016	-0.005
	(0.005)	(0.007)	(0.007)	(0.012)	(0.005)
Apr	-0.008*	-0.009	-0.006	-0.014	-0.006
	(0.005)	(0.007)	(0.006)	(0.013)	(0.005)
May	-0.003	0.001	-0.009	0.002	-0.004
	(0.005)	(0.008)	(0.006)	(0.015)	(0.005)
Jun	-0.003	-0.004	-0.003	-0.003	-0.003
	(0.006)	(0.010)	(0.007)	(0.015)	(0.007)
Jul	-0.002	-0.006	0.002	-0.008	-0.001
	(0.005)	(0.008)	(0.007)	(0.014)	(0.006)
Aug	-0.003	-0.006	0.000	-0.009	-0.002
	(0.006)	(0.008)	(0.007)	(0.015)	(0.006)
Sept	-0.004	-0.006	-0.002	-0.007	-0.002
	(0.005)	(0.008)	(0.006)	(0.014)	(0.005)
Oct	0.002	0.008	-0.006	0.005	0.001
	(0.006)	(0.009)	(0.006)	(0.014)	(0.006)
Nov	0.008	0.012	0.005	0.011	0.009
	(0.006)	(0.009)	(0.007)	(0.014)	(0.007)
Dec	0.004	0.007	-0.000	0.009	0.003
	(0.005)	(0.008)	(0.005)	(0.014)	(0.005)
Atlantic	0.006	0.006	0.008	0.001	0.009

	(0.007)	(0.009)	(0.011)	(0.014)	(0.008)
BC	0.002	-0.001	0.007	0.013	-0.000
	(0.007)	(0.010)	(0.010)	(0.023)	(0.007)
Ontario	0.002	0.004	0.000	0.004	0.001
	(0.006)	(0.009)	(0.008)	(0.013)	(0.007)
Prairies	0.002	0.004	-0.001	0.003	0.001
	(0.007)	(0.010)	(0.008)	(0.012)	(0.008)
Quebec	0.022**	0.025*	0.020	0.063***	0.012
	(0.009)	(0.014)	(0.012)	(0.025)	(0.009)
Year2016	-0.014*	-0.016	-0.011	-0.020	-0.012
	(0.007)	(0.010)	(0.010)	(0.015)	(0.008)
Year2017	-0.023***	-0.034***	-0.010	-0.059***	-0.014**
	(0.007)	(0.010)	(0.009)	(0.017)	(0.007)
Year2018	-0.034***	-0.036***	-0.033***	-0.046***	-0.030***
	(0.008)	(0.010)	(0.012)	(0.015)	(0.009)
N	27741	15194	12547	5864	21877

*** p<.01", "** p<.05", "* p<.1. Standard errors are in parentheses.

Chapter 3

Physician Labour Supply: Efficient Use of General-Purpose Surveys

Abstract

Physician labour supply trends in Canada from 1987 to 2021 are examined together with the correlates of the current perceived physician shortage. A survey weighting technique with general applicability is employed that enhances estimation quality for profession-specific analysis using the modest sample sizes of general-purpose surveys. Our findings indicate a 16.5% decline in average work hours among practicing physicians during the study period. The reduced work hours among male physicians contribute significantly to this decline. The growing representation of females within the workforce, characterized by fewer hours than their male counterparts, also contributes to the downward trend. Moreover, the aging of the physician workforce contributes noticeably to the overall decline in average work hours. Particularly, there has been an approximately 90% rise in physicians aged 65 and older between 1989 and 2021. These senior physicians work notably fewer hours than their prime-age counterparts. Increasing lengths of absences (e.g., vacations) among physicians also contributes to the decline. Though we cannot distinguish cohort effects from year effects, successive cohorts are observed to work fewer hours than preceding ones.

3.1. Introduction

Perceptions of physician shortages in Canada, reflected in long wait times for primary, specialist and emergency care, is a central concern of policymakers, researchers, and healthcare managers (Canadian Institute for Health Information (CIHI), 2022b; Chan, 2002; Islam et al., 2023; Jones, 2022; Kermode-Scott, 1999; Malko & Huckfeldt, 2017; Purdon & Palleja, 2023; Roos et al., 1998; Smart, 2022; Wait Time Alliance, 2015). Most studies quantify the levels or trends in the supply of physician labour using indirect measure such as headcounts per capita or measures derived from administrative data such as the number or value of billings (e.g., Adams et al., 2017; Baxter, 2000; Canadian Institute for Health Information, 2021; Roos et al., 1998; Vogel, 2013). Very few studies focus on physicians' hours and weeks of work (actual labour supply) because of the perceived lack of occupation-specific data (an exception is Islam, Kralj and Sweetman, 2023). To overcome this, we use Statistics Canada's Labour Force Survey (LFS; its national labour market survey that is similar to the American Current Population Survey or CPS) that identifies occupation, has a good response rate and a variety of labour supply questions. We are not the first to use a general-purpose survey to study health professionals. In the United States, Frogner & Dill (2022), Hu & Dill (2021), and Staiger et al. (2010) use the CPS to look at physicians and other health professions. However, like the CPS, there are limits to using the LFS for occupation-level studies since it has modest samples of workers in specific occupations and the general-purpose survey weights provided by statistical agencies are not well suited for profession-specific analysis. Inferences drawn from the ensuing estimates are limited.

To more efficiently estimate parameters associated with physician labour supply that take advantage of the LFS questions, we examine the feasibility of increased precision from improved survey weights. Our approach leverages the fact that, for regulated health occupations, we have high quality measures of workforce characteristics that can be employed for this purpose. Our substantive economic question focuses on physicians given current policy priorities, but in principle our approach may be employed for many other health occupations since similar auxiliary data are common given that the regulatory process provides a (normally annual) census of each occupation. We contrast unweighted estimation, estimation using Statistics Canada's general-purpose survey weights, and two alternative approaches to generating occupation-specific weights using auxiliary data from regulatory authorities. To generate occupation-specific weights, first, we use a traditional raking (i.e., Iterative Proportional Fitting or IPF) method as implemented by Kolenikov (2014, 2019); see Dever et al. (2008) for an introduction to survey weights in the statistics literature. We contrast this to a little used generalized method of moments (GMM) technique introduced by Imbens & Lancaster (1994; hereafter I&L), and Hellerstein & Imbens (1999; hereafter H&I). We observe what we interpret as a marked improvement in efficiency, and believe that this approach may allow the LFS, CPS and other similar general-purpose surveys to be employed to study health workforce issues where inference has been historically problematic.

Centrally, we investigate self-reported hours of work data from the LFS between 1987 and 2021 to identify trends in physician labour supply and the associated characteristics of these workers. This extends research by Crossley et al. (2009), Sarma et al. (2011), Islam et al. (2023), Staiger et al. (2010), and Winkelmann et al. (2020). A key puzzle is to understand why, despite Canada's stock of practicing physicians growing by about 98% from 1987 to 2021 while population growth was only about 45%, there were perceptions of a physician surplus in the late-1980s/early-1990s but ongoing reports of physician shortages in recent decades (Baxter, 2000; Chan, 2002; Jones, 2022; Smart, 2022). Indeed, in the 1990s, when physicians per capita appeared to be relatively low by the norms of the 2020s, fearing surpluses Canadian provincial governments elected to constrain medical school enrolment growth (Dauphinee, 1996; Roos et al., 1998; Ryten, 1995); this policy that is commonly, and probably partly erroneously, attributed to the Barer & Stoddart (1991) report.

Reconciling the growth in physicians per capita with increasing shortages requires understanding not only physician labour supply but also trends in population demand, particularly that driven by population aging (Islam et al., 2023). A full explanation requires an analysis of changing practice patterns and the composition of the physician workforce, which leads us to take an intensive margin approach in this paper.

We find that average work hours declined by about 16.5% or 7.9 hours per week between 1989-2021.[1] Although the proportion of male physicians in the workforce declined significantly from around 80% in 1987 to 56% in 2021, males remain the dominant gender in the physician workforce. A substantial reduction in their weekly work hours by 7 hours during the study period, combined with their significant presence in the workforce, has an enormous impact on the overall decline. The growing share of female physicians, from about 21% in 1989 to 44% in 2021, in the workforce characterized by fewer hours roughly plays the same role as male physicians in decline. These results are consistent with existing studies (Crossley et al., 2009; Sarma et al., 2011; Islam et al., 2023). Besides, we find aging of the physician workforce is an important factor in the declining trend. Our study shows that physicians aged 65 and over have increased from about 8% to 15%, about a 90% increase. This group works significantly fewer hours than its prime-age counterparts. Moreover, the rate of absences from work (lasting one

[1] The reduction of work hours reported in Chapter 1 differs from that in Chapter 3, being 20.6% in the former and 16.5% in the latter. This difference arises because Chapter 3 employs a three-year moving average whereas Chapter 1 does not. The numerical reductions would be identical if the chapters presented the data in the same way.

week or more) increased from about 9% to 11%, and the length of these absences also grew from around 4 weeks to 7 weeks, contributing to the overall decline in average work hours.

Factors such as marital or immigration status do not make a significant contribution to trends in average work hours. However, interactions of children with gender highlights a substantial gap. Male parents increase their work hours compared to male physicians without children, while female parents reduce their work hours compared to female physicians without children. Regional analysis reveals noticeable differences between provinces, especially in the early years when average work hours are high. However, convergence is observed in the later years when average work hours decrease.

Furthermore, we detect signs of cohort effects, with each subsequent cohort working fewer hours than their predecessors. However, it remains challenging to disentangle cohort effects from year effects.

The subsequent sections of this paper are structured as follows: Section 3.2 provides an overview of the data, Section 3.3 ascribes the methodology employed, Section 3.4 addresses the necessity of reweighting, Section 3.5 presents the results, and, lastly, Section 3.6 concludes the research.

3.2. Data

Statistics Canada’s monthly microdata from January 1987 to December 2021 are central to our analysis. In addition, the Canadian Institute for Health Information (CIHI) provides aggregate physician population counts and means -- by age, gender and geographic region -- derived from annual regulatory college registration censuses that are used as auxiliary data to generate survey weights. Until more recently than the end of our data, Statistics Canada collected information regarding what it termed the respondent’s “sex” as a binary male/female categorization, we believe that many in the group under study would have responded with their gender; we therefore choose to refer to male/female as identifying gender (albeit limited to two categories). However, it is not known with certainty how respondents actually self-identified in response to the Statistics Canada question and whether such self-identification might have changed over time. We start in 1987 because of LFS occupational coding issues, and end the sample in 2021 because of CIHI auxiliary data availability.[2]

3.2.1. LFS Data

Physicians' self-reported hours of work data are obtained from the LFS, a monthly household survey conducted by Statistics Canada. Besides occupation and hours of work, the LFS collects data on other employment, personal and family characteristics for civilian, non-institutionalized Canadians aged 15 and older living in a province (i.e., omitting the territories). At present, the LFS's monthly sample comprises approximately 56,000 households consisting about 100,000 observations. Practicing physicians are identified using their labour force status, National Occupational Classification (NOC) code, and North American Industry Classification System (NAICS) code. Physicians with full-time student status or aged below 28 are excluded from the analysis sample to approximate CIHI's exclusion of medical residents. No upper age limit is imposed since some physicians are known to practice into their 80's and CIHI counts these physicians in the physician population. The maximum age of any observation in our samples falls in the range 85-90. The LFS data do not contain information on the graduation year of physicians; consequently, we define cohorts according to birth year. Hours of work are capped at 100, affecting less than 1% of physicians. Details are in Appendix 1.

[2] It is common, and reasonably reliable, to project this type of regulatory data a year or two ahead, which would allow us to extend the analysis to 2023. But, we prefer to employ only measured data to generate the survey weights in this initial use of the approach in this context.

We append the monthly surveys into annual samples. The annual physician samples vary from just above 1,700 to almost 4000 observations and the pooled 1987-2021 sample has about 98,000 observations. Since the LFS is a six-month rotating panel (compared to the CPS's 4-8-4 design), a physician may appear up to six times each year and may appear in adjacent years. Consistent with Statistics Canada's apparent approach, and for the purpose of constructing weights, the repeated cross-sectional units are treated as independent. In ancillary work we observe that the LFS data cannot reject a test with the null hypothesis that attrition meets a missing completely at random (MCAR) assumption for hours of work. To satisfy Statistics Canada's data release requirements, adjacent provinces with small populations are combined; the Atlantic Provinces (Nova Scotia, New Brunswick, Newfoundland, and Prince Edward Island) form one region, and Manitoba and Saskatchewan another.

3.2.2. CIHI Data

Physician population moments and counts are obtained from Canadian Institute for Health Information (CIHI, 2022a). CIHI variables common with the LFS are counts of male and female physicians by region, the average age of physicians, and the number of physicians by age group and region.[3] These data are considered to be reliable measures of physician counts based on licensure information.

[3] Urban and rural classifications are not used since CIHI and the LFS have inconsistent definitions.

3.3. Methodology

Given our focus on a single occupation with a relatively small sample size, we evaluate potential efficiency gains from using auxiliary data to generate occupation-specific survey weights and improve efficiency as discussed by L&I and H&I. We compare statistics generated using unweighted data, Statistics Canada's generic LFS weights, and occupation-specific weights generated by a traditional raking/IPF method and the GMM approach espoused by H&I.

In econometric and statistical practice, the use of survey weights can be controversial when estimating coefficients in correctly specified regression models where the regressors are also the weighting variables, though there is no disagreement regarding the need for their use in calculating means or average marginal effects (e.g., Cameron and Trivedi 2005, 2022). In contrast, statisticians have long pointed out that when models are not correctly specified (e.g., when linear or quadratic approximations to more complex functional forms are employed) weights can change estimates appreciably (Deville & Särndal, 1992; Korn & Graubard, 1995). Further, many surveys, such as the LFS and CPS, have complex survey designs where the sampling probability varies markedly across individuals as a function of the weighting variables which, among other issues, affects the precision of estimates and in practice ignoring survey weights frequently makes estimates appear overly precise. Since much of our work focuses on simple averages for various subpopulations, weighting is clearly relevant. Of course, weights only ever solve a selection on observables, or missing at random (MAR), problem. They do not address issues arising from, for example, self-selection, non-random non-response and/or attrition.

3.3.1. The Raking or Iterative Proportional Fitting Method

Raking, also known as the Iterative Proportional Fitting (IPF), is a process that calibrates survey weights using known population/control totals/counts from auxiliary data so that the weighted data are representative of the underlying population with respect to the calibration variables employed (Kolenikov, 2014, 2019). For example, if the number of females in each physician speciality is a set of calibration variables, then for each speciality the weighted number of female physicians in the survey is set to the count totals in the population. In their simplest form calibrated weights, w_c, are defined as:

$$\sum_{n=1}^{N} w_c \, I(i \epsilon C) \; = N_c$$

where I is an indicator that takes value 1 for females in specialty C, and 0 otherwise. N_c is the total number of female physicians in specialty C in the target population, and w_c is the calibrated weight for the same group. N is the total sample size and n indexes observations. With many calibration variables for each observation, equality is not normally satisfied and the method iteratively adjusts the sample data by assigning weights to each respondent to optimize the match between the sample and population sums in each cell. For details of the matching algorithm, see Deming & Stephan (1940); Deville & Särndal (1992); and Kolenikov (2014, 2019) whose implementation we employ.

3.3.2. The Hellerstein and Imbens Method

H&I's approach employs what they term generalized method of moments (GMM), though in many applications it might be more accurately called method of moments (MM) since it is exactly identified if the model of interest is not over-identified (e.g., if the model of interest is (weighted) ordinary least squares (OLS) or probit). It simultaneously solves two sets of equations. The first set is derived from the model of interest (e.g., OLS or probit, but potentially a wide range of models) and comprises weighted normal equations or weighted score functions. The second set of equations comes from constraints imposed on the weighted sample moments using the auxiliary (population) data (e.g., uncentered first, second, and cross moments). One practical advantage of GMM compared to IPF is that it more easily accommodates a wider range of auxiliary data constraint types. IPF most easily accommodates cell counts, whereas H&I can, additionally, easily accommodate means and higher moments of continuous variables and interactions between continuous and discrete variables.

H&I present a 1-step procedure where GMM solves both sets of equations simultaneously and returns weighted estimates from the regression of interest. However, in Appendix 2, we prove that in many common cases a numerically identical 2-step approach may be employed. A first step can be used to obtain the weights, and then the weights may be used in a second step to obtain the (same) estimates of interest. Splitting the problem into two steps frequently makes estimation easier and much faster, and it allows a single set of weights to be employed in multiple regressions of interest, and in generating summary statistics, without re-estimating the weights each time. This makes the approach much more practical.

Following H&I, in the OLS context, GMM estimates weighted least squares (WLS – i.e., weighted OLS) to obtain β coefficients as:

$$\widehat{\beta_{WLS}} = [\textstyle\sum_{n=1}^{N} \widehat{w_n}\, x_n x_n']^{-1} [\textstyle\sum_{n=1}^{N} \widehat{w_n}\, x_n y_n] \qquad (1).$$

The weights, $\widehat{w_n}$, solve

$$max_w \textstyle\sum_{n=1}^{N} ln\, w_n, \text{ subject to } \textstyle\sum_{n=1}^{N} w_n = 1 \;\; \& \; \textstyle\sum_{n=1}^{N} w_n \,.\, h(y_n, x_n) = 0 \qquad (2),$$

where *h(.,.)* is a moment restriction, for example the difference between the age in the sample and the mean age in the auxiliary data (i.e., the population) which, when appropriate weights are selected, should equal zero.

Focusing first on H&I's original 1-step approach, equations (1) and (2) can be written within a GMM framework as:

$$E[\psi(y,x,\beta,\lambda)] = E\begin{bmatrix}\psi_1(y,x,\beta,\lambda)\\ \psi_2(y,x,\lambda)\end{bmatrix} = E\begin{bmatrix}x(y-\beta' x)/\left(1+e^{\lambda' h(y,x)}\right)\\ h_{k(y,x)}/\left(1+e^{\lambda' h(y,x)}\right)\end{bmatrix} = 0 \qquad (3).$$

The exponential, e, in equation (3) avoids negative weights. In minimizing its weighted criterion function, GMM simultaneously chooses β, and also λ where the latter makes the weighted sample moments close to those in the population. For example, in $\psi_2(y,x,\lambda)$, GMM sets $E\left[h_k(y,x)/\left(1+e^{\lambda' h(y,x)}\right)\right] = \mathrm{E}[\mathrm{w_n h_k(y_n, x_n)}] = 0$ with $w_n = 1/\left(1+e^{\lambda' h(y_n, x_n)}\right)$, where k indexes calibration/auxiliary variables. If k = age and we focus on the first moment, then $h_k(y,x)$ could be $h_{age_n}(y_n,x_n) = Age_n - \overline{Age_{pop}}$; where $n = 1, \dots \dots, N$, Age_n is age of individual n, and $\overline{Age_{pop}}$ is the mean value of age in the population. The expected value of $h_{age_n}(y_n,x_n)$ in a representative sample is, obviously, zero.

Parts of our analysis estimates simple weighted averages, especially the average hours of work of physicians in annual samples and sociodemographic subsamples. The counterpart of equation (1) is then:

$$\widehat{\mu_{\mathrm{WLS}}} = \left[\sum_{n=1}^{N} \widehat{w_n}\right]^{-1}\left[\sum_{n=1}^{N} \widehat{w_n}\, y_n\right] \qquad (4),$$

which is equivalent to WLS regression with only an intercept. The weights, $\widehat{w_n}$, are an implementation of equation (2). Under GMM:

$$E[\psi(y,x,\beta,\lambda)] = E\begin{bmatrix}\psi_1(y,x,\lambda)\\ \psi_2(y,x,\lambda)\end{bmatrix} = E\begin{bmatrix}(y-\mu_y)/\left(1+e^{\lambda' h(y,x)}\right)\\ h_{k(y,x)}/\left(1+e^{\lambda' h(y,x)}\right)\end{bmatrix} = 0. \qquad (5)$$

Equation (5) returns weighted sample averages in one-step.

However, for simple means, and more generally for just-identified WLS with regressors, this is numerically equivalent a two-step approach. In many contexts, the two-step approach is much less expensive in terms of computer and researcher time, which makes the H&I method much more practical than it would be if the one-step approach were required to, for example, repeatedly solve the full GMM problem to generate a simple descriptive statistic for each variable. The savings follows from, first, only needing to estimate the weights once for a given sample before employing them repeatedly for various second steps (e.g., in calculating each regressor's weighted mean) and, second, from the time needed for the GMM to converge increasing more than linearly in the number of parameters estimated. The two-step method first generates weights by estimating the parameters associated with $\psi_2(y,x,\lambda)$, and then employing the ensuing weights to solve $\psi_1(y,x,\lambda)$. In the second step, it is time saving to employ the software's numerically equivalent built-in WLS rather than GMM. An analytical proof is provided in Appendix 2.

More broadly, if the second step is a just-identified non-linear model (e.g., a logit or probit) then using the standard command in the second stage does not provide numerically identical results to GMM. But, if the GMM and whatever solution method is employed for the non-linear estimator are converging to the same fixed point, then the difference between the two will grow arbitrarily small in expectation as the sample size increases. Our simulations show that the one- and two-step approaches provide very similar estimates for the coefficients of interest. For clarity, we only establish the one- and two-step equivalence for just-identified models.

3.4. Reweighting the LFS Data Using the Regulatory Data

3.4.1. Comparison of the LFS to CIHI Data

Reweighting is a selection on observables (missing at random conditional on X's) approach to addressing sample data that is not representative of the underlying population because of the survey design and/or survey non-response. However, some researchers prefer to employ unweighted data, especially when the regressors employed include the weighting variables and the model is correctly specified, since poor quality estimates for the weights may increase the mean squared error of the estimates of interest (Cameron & Trivedi, 2005, 2022). However, there is no controversy about the need for such weighs when estimating simple descriptive statistics such as the group/cell means that comprise much of this paper.

Although not a test of each method, some sense of the magnitude of any improvements in estimates from using the regulatory data are examined in this section. In Figure 3.1, we plot the average age of physicians by year from CIHI's administrative files comparing unweighted estimates with weighted estimates using LFS, IPF and GMM weights. To avoid crowding the plot, 95% confidence intervals are displayed for (only) the LFS- and GMM-weighted estimates. The weighting variables for the IPF method are cell counts in each of the five age categories for each of the six regions in addition to counts of males and females in Canada and in each region. (Regions comprise large provinces, or groups of adjacent less populous provinces.) For illustrative purposes, since GMM easily accommodates both cell counts and moments of continuous variables, for the GMM approach, we employ counts in four of the five age categories in each region and another age category (30 or below) for three regions plus the overall average age of physicians, and the proportions of females in Canada and five of the six regions. Using counts in all five age groups for all regions plus the average is not feasible for

GMM because of collinearity. For the same reason, we cannot use the proportions of females for all regions and Canada together. These are the maximum number of constraints we can impose on the LFS data using available information from CIHI. We apply this set of constraints in our analysis of the sample data. To ensure that the H&I method improves estimates reasonably under this setting, we conduct a set of Monte Carlo simulations in the following subsection using the same setting and constraints. A comparison of alternative specifications of the weighting variables in the GMM context is provided in Appendix 3.4.

In almost every year, the average age of physicians is underestimated by both the unweighted and LFS-weighted estimates relative to the "truth" observed in the CIHI data. The unweighted and LFS-weighted estimates are similar, and their confidence intervals do not include the CIHI provided means in almost all years. In contrast, the IPF and GMM-weighted estimates are very similar to the CIHI estimates; they are difficult to distinguish visually on the plot, and these two series' confidence intervals contain the CIHI's means in every year. Note that, in line with the theoretical presentation regarding precision by L&I and I&H, the year-to-year variation of the average age profile is also much diminished when using the CIHI auxiliary/population data to reweight the LFS data by either the IPF or GMM method. The reweighting improves the precision of the estimates appreciably for these sample sizes. This is a non-trivial benefit.

Reasons for the LFS underestimation are examined visually in Figure A3.1 of Appendix 3.3.1. We find the LFS persistently both over-samples younger physicians and under-samples older ones (the left panel), and the LFS weights do not address this sufficiently.

In Figure A3.2 of Appendix 3.3.1, we similarly investigate gender to show that the LFS-weighted estimates of the percent female deviate markedly from the mean in the administrative data. A t-test of equality between the CIHI total and the LFS-weighted total rejects the null with a t-statistic of 5.54. Again, the GMM- and IPF-weighted estimates are very close to the CIHI percentage (almost visually indistinguishable) and we cannot reject the null that they are the same. For both, t-statistics of the tests similar to that for the LFS weights are close to zero. Overall, the LFS weights are not generated for physician-specific (or any sub-group specific) analysis and estimates can be improved by reweighting using regulatory/auxiliary data.

3.4.2. Monte Carlo Simulations Contrasting the GMM and IPF Methods

Given that the LFS data would benefit from improved survey weights and that both IPF and GMM are candidates, we conduct Monte Carlo simulations looking at weekly hours of work to compare them. To mimic our context's correlations among the regressors, the simulation design combines the annual LFS physician samples from 1987-2020 into a pooled dataset that we take as the "true" population.[4] From this population, small simulation samples are selected with sampling probabilities based on gender and geography that are detailed in Appendix 3.2. We employ different sample sizes, but emphasize 1700 since it is approximately the lower bound of the LFS annual data sample size.

If we use the same information (i.e., same specification of the same weighting variables) for these two different approaches, the performance of IPF and GMM is essentially identical in that they have virtually identical means squared errors (MSEs). The results can be seen in Appendix 3.3.3 for samples of 1700 (Figure A3.3).

However, one virtue of the GMM approach is that it not only allows matching on cell counts, which is all that is feasible in the IPF approach, but it also very easily allows for the use of means, uncentered higher order moments and cross-products of essentially unrestricted (except

[4] In constructing population for the Monte Carlo simulation, we use data between years 1987-2020, as more recent data were not available when we started the simulation.

by the sample size) combinations of continuous and indicator variables. Therefore, for a sample of size 1700, Figure 3.2 illustrates the distributions of the incremental value, in estimating mean hours of work, of incorporating the mean of the continuous variable age with four of the five age categories in each region and other category (below 30) for four regions as weighting variables in addition to the gender and region variables as described in the appendix 3.3.2.

As expected, given the non-random sampling design, the distribution of the unweighted simulation estimates are centered around a point far left to the true mean value, which is presented by the vertical solid line in each panel. Both GMM and IPF approaches to weighting effectively center the distribution around the true value and have very similar distributions. However, incorporating continuous variable age allows the GMM estimator to have a slightly smaller MSE. In general, adding additional relevant constraints, even higher order terms for already included variables, appears to improve the estimates. The choice between the two methods for generating occupation-specific survey weights appears to be primarily one of researcher convenience and whether it is feasible to easily employ additional restrictions in the GMM context beyond what is feasible for the IPF raking method. The flexibility of the GMM approach on this front is a practical advantage. Increasing or decreasing sample size does not alter these results as plotted in Figure A3.4 for 3500 and Figure A3.5 for 850 in Appendix 3.3.3. We next explore how important this might be in practice.

3.4.3. Out of Sample Prediction from Reweighted Sample Estimates

It is difficult to do a predictive (out of sample) validation/comparison of the two approaches in our application since we exhaust the CIHI data in defining the weighting variables; that is, there is very little unused “truth” in the CIHI data against which we can compare weighted predictions since variables in the auxiliary data are all fundamental and all moments are employed in generating the weights. However, we contrive to create a situation that allows for such a validation/comparison. We employ the age group “60 and above” to generate survey weights, and then try to use target the number of physicians aged 65-74 using correlations among the other weighting variables for identification. We know the population count of those aged 65-74 in the CIHI data for comparison. Although we do not show it here, using the same restrictions the IPF and GMM methods have remarkably similar results. In this simulation we illustrate the value of the same single piece of information as in section 3.2; we also match on the mean of the continuous variable age. We compare MSEs in a simulation using unweighted, and LFS-, IPF-, and GMM-weighted estimates of the share of physicians who are aged 65-74.

Table 3.1 represents the results, which shows the GMM with mean age in the constraints, has the lowest MSE and, oddly, the LFS-weights perform more poorly than the unweighted estimates. This exercise shows that the expanded set of constraints available under GMM can be helpful in reducing the MSE. We employ the GMM method throughout the remainder of this paper.

3.5. Results

Turning to the paper’s substantive economic question, we investigate the LFS data applying GMM weights to improve the precision of our estimates. We examine hours of work over time, across demographic groups, and by birth cohort. Beyond descriptive regressions, counterfactuals predictions are presented to get a sense of the determinants of the observed aggregate patterns in the data.

3.5.1. Trend Analysis

Focusing on annualized trends from 1987 to 2021 (or 1989 to 2021 since we use a 3-year moving average), the left-hand side of Figure 3.3 shows average hours of work by all physicians. Two measures are produced: one excluding those who report a full-week absence (e.g., a week of vacation), and one including such physicians with their weekly hours recorded as zero.[5] The average is markedly affected by the treatment of this group, and one advantage of the LFS compared to the occasional historical surveys of hours of work conducted by the Canadian Medical Association (CMA, 2019) and studied by Crossley et al. (2009), and Sarma et al. (2011), is that this feature can be captured. These CMA surveys have almost always asked a retrospective question about average hours worked per week in the year, which (by comparison to these estimates) appears to have been interpreted as hours worked in a "usual" week when the physician had positive hours.

As can be seen in Appendix Figure A3.6, the rate of absences has been increasing since about 2005, while the length of observed absences increased over the entire period – average from about 4 weeks to over 7. In the remainder of the paper, we include practicing physicians who report zero hours in the week since changes in, for example, the frequency and duration of weeks of holidays is an element of total labour supply. However, with or without including absences, Figure 3.3 shows a downward trend in hours with more than 15% between years 1989-2021. However, the speed of the reduction is much higher before 2005 than following it.

The right-hand side of Figure 3.3 presents trends across selected percentiles of the distribution of hours. Substantial dispersion, and a long tail at lower hours (e.g., the gap between

[5] Well-known selection issues arise in the measurement of the length of absences in a survey that focusses on a single week each month. Especially, long duration absences are more likely to be observed. Nevertheless, these trends are informative since the sampling method has been stable.

the 10th and 25th percentiles is larger than that between the 75th and 90th) are easily observed. Overall, while there are some modest differences in timing and degree at different quantiles, the decline is observed across the entire distribution.

Consistent with previous observations, Figure 3.4 shows that female physicians on average work fewer hours than their male counterparts, although males' average hours of work declined much more over these years. During the same period, the proportion of female physicians increased from around 21% to 44%. Aggregate labour supply was therefore affected by both the decreasing hours of males and the increasing share of female physicians with lower average hours.

Our data also support investigating age and birth cohort effects. As seen in Figure 3.5, age subgroups between ages 28-64 supply similar average number of hours. We display the pointwise 95% confidence interval around the subgroup 28-39 (suppressing other confidence intervals for expositional purposes), to illustrate that most other means (except for the 60-64 age group in some months) fall within it. However, physicians aged 65 and older supply fewer hours than their prime-age counterparts; however, in contrast, their hours have not fallen over time. A key issue for aggregate labour supply is that the proportion of physicians aged 65 and over has increased from about 8% in 1989 to 15% in 2021. This is a contributor to the decline seen in Figure 3.3.

As seen in Figure 3.6, compared to physicians without children in the household, the presence of children, especially children aged 5 and younger, is associated with large differences in labour supply. But the effects are in opposite directions for females and males. Female and male physicians without children work very similar hours and exhibit similar trends over time in hours of work. In contrast, female physicians with young children work fewer hours while males with children work more. A similar comparison shows that the presence of older children (ages 6 to 24) exhibits a much smaller hours reduction for females, but the impact for males is comparable to that for younger children.

In terms of trends over time, for females there is no evidence of a trend in average hours of work across the entire period under study for those with children in either age group. However, there is a decline in hours of work for females with no children in the household between 1989 and about 2005, after which hours of work flattened for this demographic group; after 2005 there is effectively no gap in average hours between females without children and those with older children in the household. For males, the trends over time are quite different than those for females. The decline in average hours is much larger among those with children, of any age, in the household compared to those without children. At the end of the period the “presence of young children” hours premium – relative to males with no children and their household – declined from approximately 10 hours a week to about five hours a week.

Regional differences in hours of work are explored in Figure 3.7. To illustrate estimation uncertainty without making the plot unreadable, we display a 95% CI around physicians’ hours of work in Ontario, which has the smallest confidence region. Overall, there is convergence in hours of work with the variance across regions being much larger at the start than near the end of the period. Although all regions show a decline, the decline is greater for regions with higher hours initially.

Neither marital status nor the same interacted with the presence of children are associated with differences in hours. Similarly, we observed no notable hours differences between immigrant and non-immigrant physicians. In both cases, the results are not presented to save space.

3.5.2. Birth Cohort Analysis

Birth cohorts are tracked over time in Figure 3.8 for males and 3.9 for females. Early female cohorts cannot be presented due to small cell sizes. To facilitate interpretation, we label the years 1995 and 2005 for relevant birth cohort age profiles, and age 30 and 45 for the year profiles. We calculate the average age by applying GMM weights to the sample data of each cohort in each year. In the plot, ages 30 and 45 are approximations since averages are not round figures. For older cohorts, e.g., Cohort 1, we neither observe age 30 nor 45; for some other cohorts, we observe either of them; therefore, on the plot, we can see these labels are missing for some cohorts. It is not possible to identify cohort effects separately from common year effects that affect all birth cohorts approximately simultaneously in these simple plots, but they are nevertheless instructive.

Reflective of the trends by gender in Figures 3.4 and 3.6, the gaps across male birth cohorts are larger than those among their female counterparts. Indeed, there is not much evidence of consistent hours gaps across birth cohorts or years for females. For males, there are clear gaps across birth cohorts both as they age and over time, but these are descriptive and do not identify birth cohort effects. To verify that these gaps are statistically significant, we perform pairwise t-tests on the mean difference of hours of work at common ages for overlapping portions of cohort trajectories, which tells us that hours of work at the common age range are statistically different across birth cohorts. For example, between ages 51-60, we observe three cohorts overlap in our data, e.g., birth years between 1931-1940, 1941-1950, and 1951-1960. Test results for these cohorts show that their hours of work are statistically significantly different from each other at the 5% level. Of course, this does not imply that there are fixed birth cohort differences in average hours give the declining trend on hours over time.

We depict cohort-specific hours of work combined for males and females in Appendix Figure A3.7. It shows the patterns of combined trends are dominated by the patterns of male cohorts as seen in Figure 3.8.

3.5.3. Regression Analysis

Table 3.2 presents ordinary least squares (OLS) regression results across selected time intervals. The "1987-2021" column includes all observations from 1987-2021, and the columns to the right are limited to subsets of those years where we first split the full period into two, and then divide the second half into additional 2 subperiods. In our estimation, we employ the following weighted regression model:

$$Hours_{i,t} = \beta_0 + \beta_1 Fem + \sum_k \beta_k Age_k + \sum_c \beta_c Child_c + \eta_1 Hosp + \eta_2 Mar + \sum_k \delta_k Age_k * Fem + \sum_c \delta_c Child_c * Fem + \sum_r \delta_r R_r + \sum_t \theta_t Y_t + \sum_m \gamma_m M_m + \varepsilon_{i,t} \quad (6),$$

where: $Hours_{i,t}$ is the hours of work of physician i in year t; Fem is an indicator variable for females; Age_k, a set of indicator variables for k=4 age groups, i.e., 46-49, 50-59, 60-69, and 65 & Above, with age group 28-39 omitted in the intercept; $Child_c$, is a set of indicators for the presence of children with c=2 age groups, i.e., children's age <= 5 years and 5 <age <=24, with no children as the omitted group; Hosp is an indicator for practice setting, i.e., in hospital or community; Mar is an indicator for marital status, i.e., common law/married or other; R_r, is a set of regional dummies with r representing the number of regions, omitting Alberta in the intercept; Y_t is a set of year dummies with t number of years; and M_m is a set of month indicators to capture seasonality, with m number of months omitting January in the intercept. The error term, ε, is clustered at the physician level given that the rotating panel design implies that physicians may be repeated up to six times. All Greek letters represents the coefficients of the respective variables.

In Table 3.2, the coefficients on female suggests that those less than age 40 with no children worked fewer hours per week than their male counterparts in the full period and all subperiods. This conditional gap is larger than what was apparent in Figure 3.6; controlling for age and the presence of children drives a large part of this difference. There is little other difference from males in hours of work across the age distribution. But overall, hours are seen to decline with age – quite modestly before age 60, and more quickly with increasing age beyond that.

Marital status shows no differences across most subgroups in terms of work hours. The presence of children in the household, especially young children, is associated with a reduction in females' work hours. Conversely, the work hours of male physicians tend to increase upon becoming parents, resulting in notable gender gaps. These observations support the unconditional analysis presented in Figure 3.6. The practice setting's influence on work hours is not statistically different from zero. Despite the absence of gaps arising from practice setting and marital status, regional variations are noticeable, mirroring the patterns illustrated in Figure 3.7.

Table 3.3 represents the analysis for the rate of absences (i.e., zero hours of work in the survey week) and the length of such absences measured in weeks. We use following probit model for the rate of absences:

$$Pr(y = 1|z) = \Phi(z) \quad (7),$$

where Φ is the cumulative distribution function of the standard normal distribution, with

$$z_{i,t} = \beta_0 + \beta_1 Fem + \sum_k \beta_k Age_k + \sum_c \beta_c Child_c + \sum_k \delta_k Age_k * Fem + \sum_c \delta_c Child_c * Fem + \eta_1 Hosp + \eta_2 Mar + \sum_r \eta_r R_r + \sum_t \theta_t Y_t + \sum_m \gamma_m M_m \quad (8),$$

where y is 1 if a physician is absent one or more weeks from work, zero otherwise. In Table 3.3, we present the average marginal effects. This implies that the effects resulting from interactions are accounted for within the original variables and, thus, are not explicitly listed.

For the length of absences, we use the negative binomial (NB) regression model with:

$$log\left(E\left(W_{i,t}\middle|Z\right)\right) = \beta_0 + \omega' Z \quad (9),$$

where

$$\beta_0 + \omega' Z \;=\; \beta_0 + \beta_1 Fem + \sum_k \beta_k Age_k + \sum_c \beta_c Child_c + \sum_k \delta_k Age_k * Fem + \sum_c \delta_c Child_c * Fem + \eta_1 Hosp + \eta_2 Mar + \sum_r \eta_r R_r + \sum_t \theta_t Y_t + \sum_m \gamma_m M_m \quad (10).$$

The dependent variable follows the Poisson distribution as under:

$$y_{i,t} \sim Poission(\mu_{i,t}) \quad (11),$$

with

$$\mu_{i,t} = e^{(\beta_0 + \omega' Z \; + offset_{i,t} + V_i,t)} \quad (12),$$

and

$$e^{V_{i,t}} \sim Gamma\left(\frac{1}{\alpha}, \alpha\right) \quad (13).$$

In the above regression, $W_{i,t}$ $is\ the\ number\ of\ weeks\ absent\ from\ work.$ Covariates are the same as in equation (6). In equation (13), α is the overdispersion parameter that governs variance. In our estimation, we use dispersion equal to 1+ $e^{(\beta_0+\omega' Z\ +offset_{i,t}+V_i,t)}$. Results presented in Table 3.3 are average marginal effects; therefore, interaction terms are not there.

In Table 3.3, the "Likelihood of Absences" column shows that female physicians are more likely to be absent from work than their male counterparts. Furthermore, as physicians age, the probability of an absence increases. Having young children (5 years or younger) also increases the likelihood of being absent from work. However, the probability of being absent decreases as the age of children increases. Practice setting and marital status neither increase nor decrease the likelihood of absences significantly. Regional variation is only noticeable for Manitoba and Saskatchewan exhibiting a lower likelihood of absence, while Alberta in the reference category.

Comparable patterns emerge in the "Length of Absences" category, where female physicians exhibit lengthier periods of leave. Furthermore, the duration of absence extends as physicians age, with parenthood also associated with lengthier leaves. Notably, no other regressor is individually statistically significant in this context.

3.5.4. Counterfactual Calculations

To quantify the magnitudes of various factors to the overall decline in hours of work, using the GMM survey weights a number of counterfactuals imposing various constraints are estimated and plotted in Figure 3.10. These counterfactuals consider various explanatory factors, such as the workforce aging and the increasing share of physicians who are female, thereby allowing the aggregate (workforce average) size and timing of various correlates to be quantified. The sample for the counterfactuals includes practicing physicians with zero hours in the survey week. (See

Appendix 4.2 for methodological details.) The observed decline in physician hours of work is depicted in the thick solid line that is the lowest on the plot in the later years, although it is not the lowest in the earlier years. At the start of the period the (three-year moving average) average weekly hours of work per physician was 47.9, and by 2021 it had fallen to 40.0. Holding the size of the workforce constant, one interpretation is that this 7.9 hours per week decline is that it is approximately equivalent to a 16.5 percent reduction in the workforce.

Prediction 1 focuses only on the aging of the physician workforce. It disentangles the evolving share of practicing physicians in each of five age categories (cells) from changing hours for each category. Counterfactually, it holds the share of physicians in each cell constant, but it allows mean hours in each to vary as observed in the data. Each annual prediction is therefore a weighted average of the observed annual mean hours for each age group, where the weights are the (three-year moving average) 1989 shares of physicians in each age group. The age-group shares and annual mean hours are estimated using the GMM survey weights. We do not distinguish between males and females in each cell, so the proportion female implicitly evolves as in the data. As seen in Figure 3.10, prediction 1's plotline does not deviate from the observed trend until about 2005, and the gap between the two only becomes economically important around 2015. Overall, of the 7.9 hour decline in average hours, the aging of the physician workforce is responsible for about 1 hour per week per physician between 2015 and 2021.

Prediction 2 is similar to prediction 1, but separates the age groups by gender. It counterfactually holds the share of each of male and female physicians, separately, in each of five age categories constant at their 1989 level, but allows mean hours in each age-gender cell to vary as in the data. This effectively age and gender adjusts the series, without constraining hours for each age-gender cell. The impact of the increasing share of female physicians, who on average work fewer hours than males, is shown in Figure 3.10 as the gap between prediction 2 and prediction 1. After growing in the early 1990's, the gap is fairly stable until about 2005 after which it increases markedly. In part, the stability until 2005 is attributable to the decline in average hours for males seen in Figure 3.4, which has an inflection point around 2005. If annual average male hours had remained stable at its 1989 level, the increasing share of physicians who are female would have had a larger impact in this scenario. Predictions 2 shows that the combination of workforce aging and the increasing share of females explains about 2.6 hours of the 7.9 hours decline.

Prediction 3 addresses a question that is conceptually distinct from the earlier two. Within each age group it counterfactually freezes male hours instead of adjusting workforce shares. It asks how aggregate hours would have evolved if average male hours of work, and only that average, remained unchanged. This is done by holding the average hours of work of male physicians unchanged at its 1989 level, but allowing the share of males/females, and female hours, to evolve as they do in the data. Prediction 3 shows average hours would have been lower in the 1990s under this assumption, reflecting the increase in male average hours of work before 1995 seen in Figure 3.4, but higher after about the year 2000. By 2021, the decline in male hours accounts for a similar share of the aggregate decline as does the increasing share of female physicians and workforce aging combined.

In Prediction 4 extends prediction 3 by, within each age group, holding both male and female hours fixed and allowing the gender composition of the physician workforce to evolve as in the data. That is, it allows for the increasing share of physicians who are female, and for gender-specific workforce aging, but does holds hours of work in each gender-age cell fixed. This is what naïve predictions in an earlier year might have estimated given an understanding of workforce aging and the changing gender composition of medical school enrolment, but not anticipating that physicians might endogenously reduce their hours of work. By 2021, this is associated with a 4.6 hour gap relative to observed average hours and illustrates how much of the labour supply decline is a “pure” hours (by gender) phenomenon.

Finally, prediction 5 is the complement to prediction 3, showing the decline exclusively attributable to falling female hours of work. It holds the gender shares and males’ hours constant within age groups, but allows annual female hours within age groups to vary. Consistent with the evidence in Figure 3.4, the effect is modest but nontrivial; by 2021 the decline in female average hours of work is of a magnitude similar to that for workforce aging, but most of the effect occurs very early in the period under study.

3.6. Conclusion

Our research provides insights regarding trends in physician labor supply in Canada between 1987 and 2021, providing some explanations for the current perceived physician shortage. Additionally, we illustrate an application of weighting techniques to improve estimation quality using a general-purpose survey for profession-specific analysis that could have wide applicability. We observe a significant decline in average work hours, amounting to approximately 7.9 hours per week, which can be translated as a 16.5% reduction in the overall physician workforce. The reduced working hours of male physicians play a pivotal role in this aggregate decline. Despite a considerable decrease in the proportion of male physicians in the workforce from roughly 80% in 1987 to 55% in 2021, the majority of physicians are still males. Their substantial decrease in work hours by 7 hours per week during the study period, interacting with their significant representation in the workforce, has influenced the overall decline substantially.

The effect of the increasing share of female physicians is also evident, with female physicians working fewer hours (around 5 per week) than their male counterparts. The impact of the increased share of female physicians is roughly the same as that of declining hours of male physicians. The rise in the rate and duration of absences also contributed to the overall decline in average work hours.

Being parents, female physicians work fewer hours than females without children, which recover as their children get old. However, male physicians increase their work hours once they have children, irrespective of the age of the children, compared to those males without children. Overall, the increase in males was offset by the decrease in females. Factors such as marriage or immigration status do not significantly influence average work hours.

Aging of physician workforce also an important factor in the declining trends, as our analysis shows that the proportion of physicians aged 65 and above has increased from about 8% in 1989 to 15% in 2021 with a significantly fewer average hours of work than their prime-age counterparts. The regional analysis reveals variations between provinces, particularly in the early years when average work hours were higher. However, convergence is observed in later years when average work hours decreased.

Furthermore, though we cannot identify birth cohort effects because time trends are confounded with them, we find signs of cohort effect. Our analysis shows that the gaps in work hours between distant cohorts are larger than between adjacent cohorts, with subsequent cohorts working fewer hours. This analysis is instructive in the policymaking to maintain a sustainable physician labour supply for Canadians.

In summary, our findings provide a comprehensive understanding of the complexities affecting physician labor supply in Canada. This research can aid policymakers, researchers, and healthcare managers in addressing physician shortages and optimizing healthcare delivery.

3.7. References

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Figures and Tables

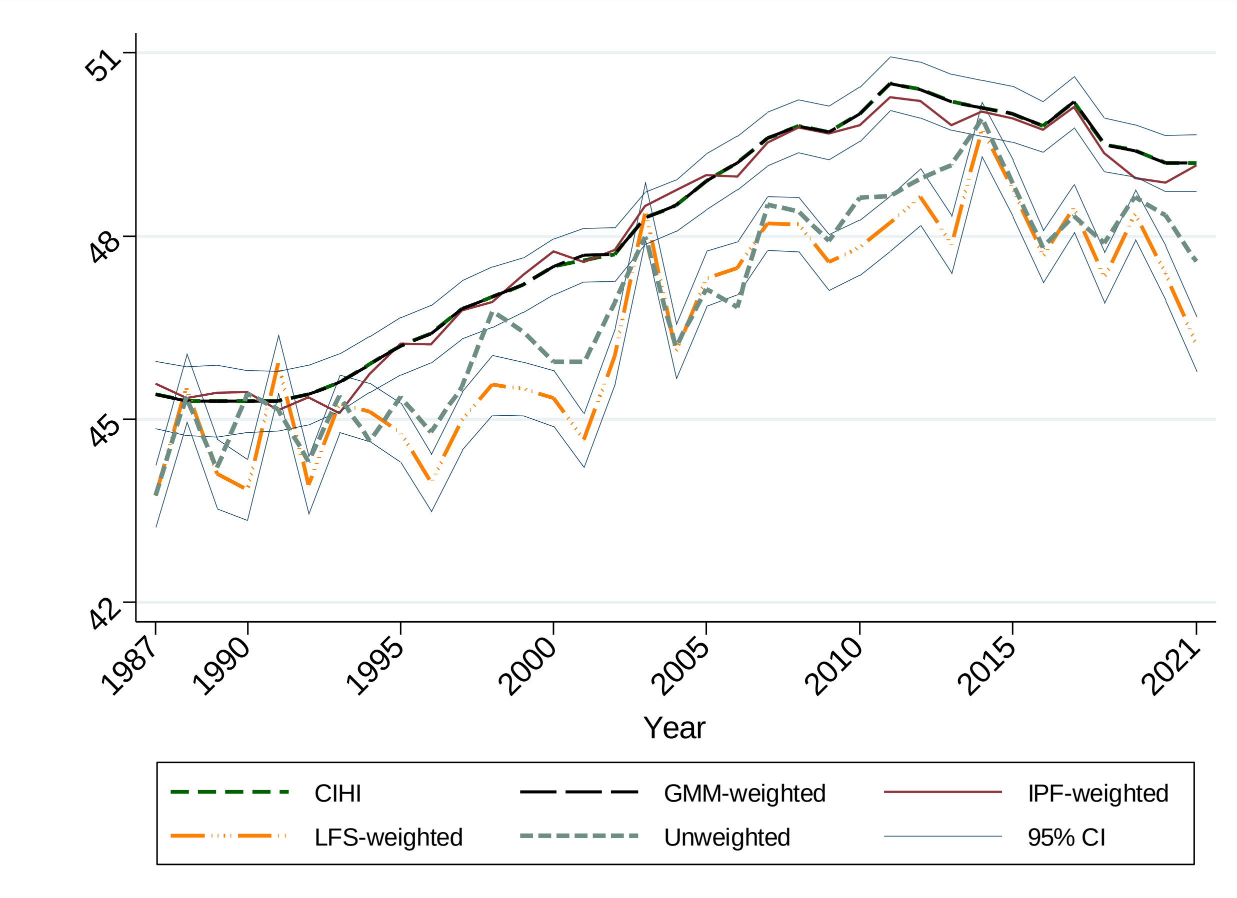
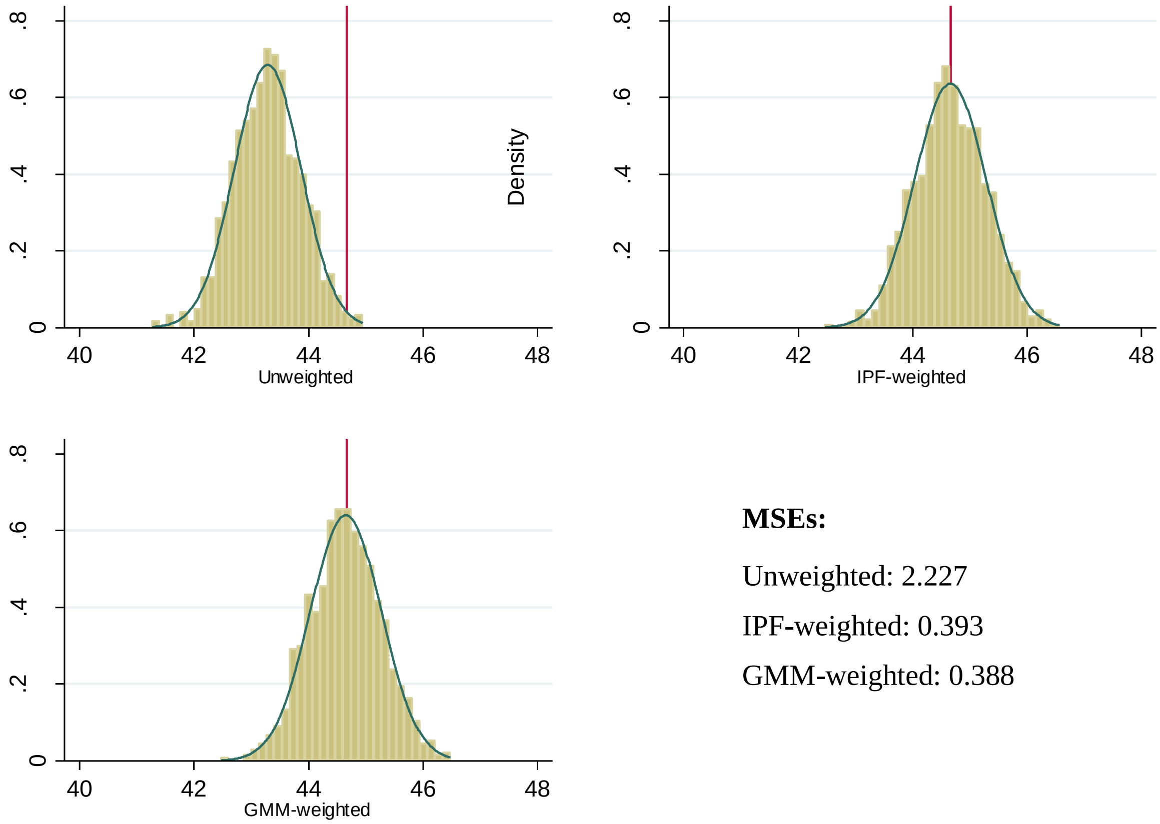


Figure 3.1: Comparison of Average Age of Physicians Using Various Methods to CIHI's Administrative Data.

Sources: Authors' Calculation using the LFS and CIHI's Public Use File.



MSEs:

Unweighted: 2.227

IPF-weighted: 0.393

GMM-weighted: 0.388

Figure 3.2: Distribution of Weighted Average Hours in MCS Using the Sample Size of 1700 with Mixed Continuous and Discrete Constraints for the GMM

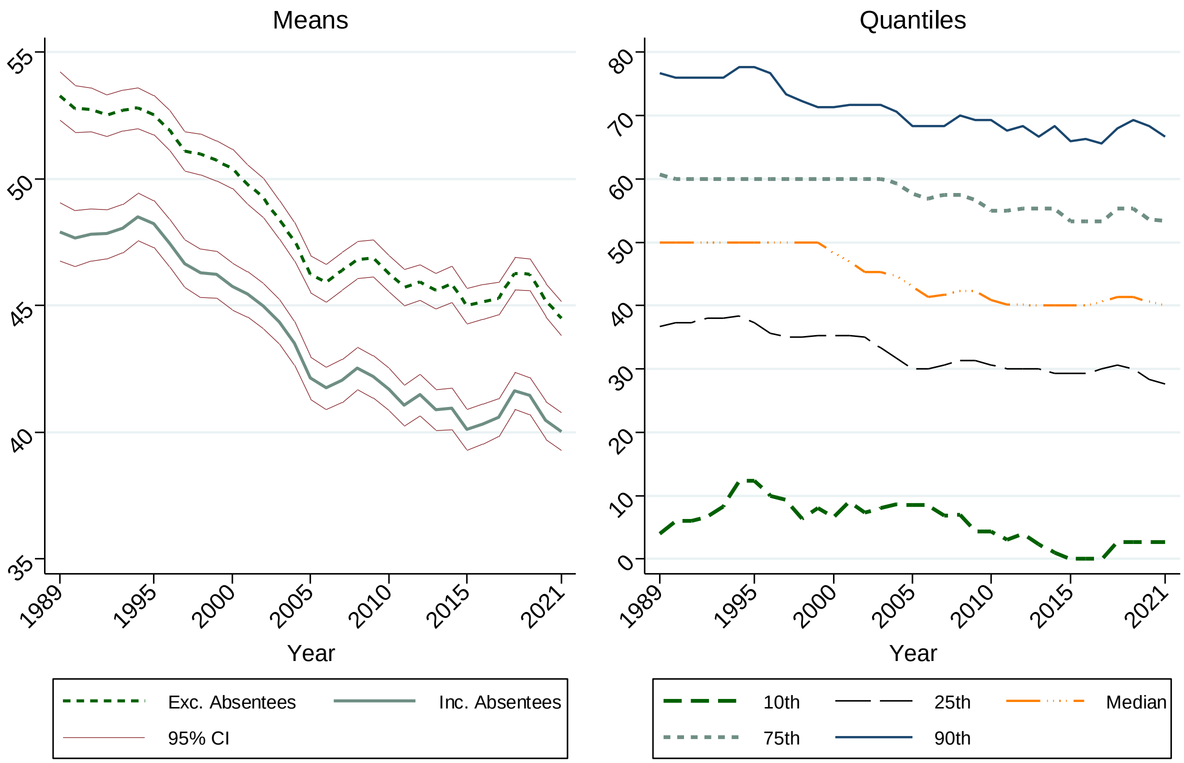


Figure 3.3: 3-year Moving Average Weekly Hours of Work Excluding and Including Absentees, and Across Quantiles.

Source: Authors' Calculation using the LFS.

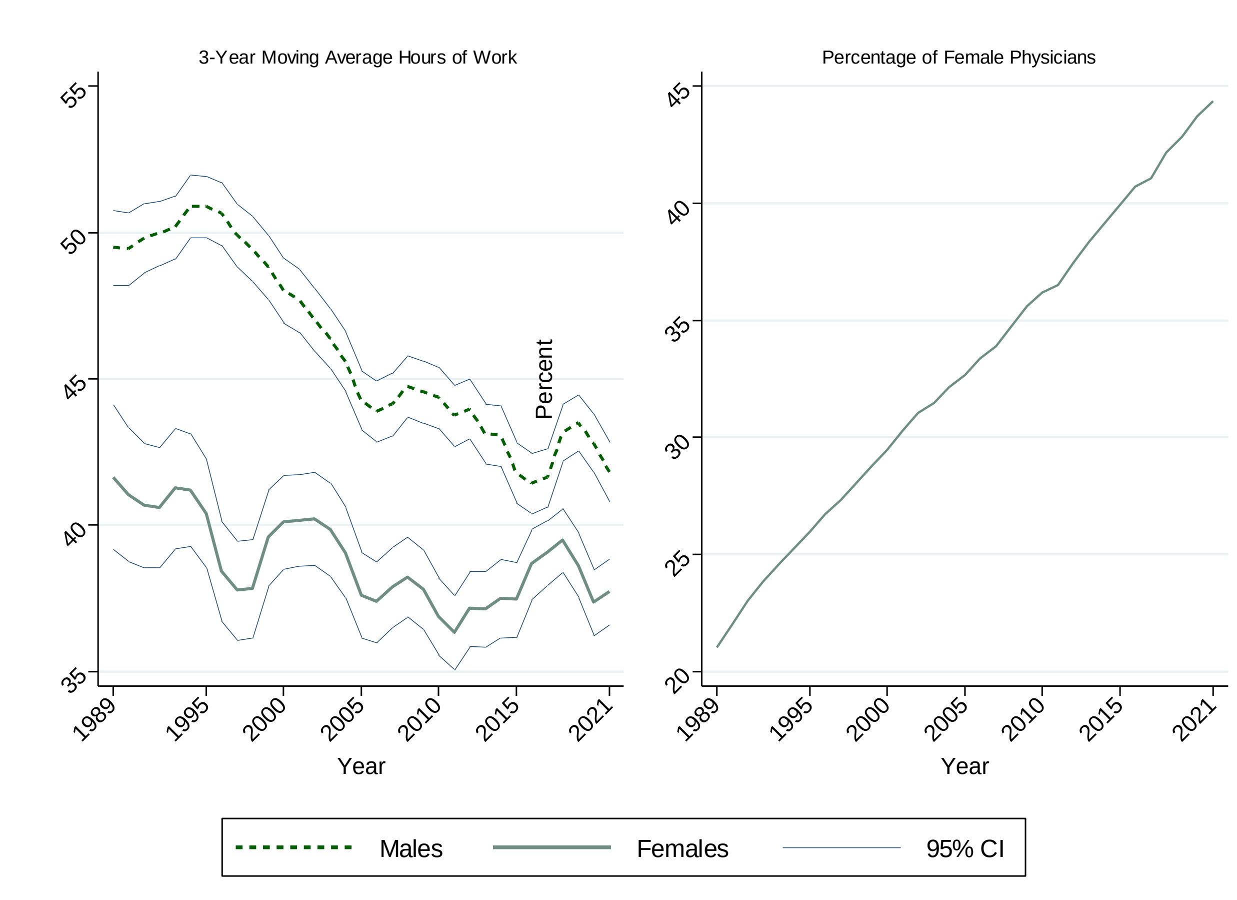


Figure 3.4: Weekly Hours of Work by Gender, and the Percentage of Female Physicians.

Sources: Left Panel — Authors' Calculation using the LFS, **Right Panel** — Canadian Institute for Health Information.

Figure 3.5: Weekly Hours of Work by Age, and the Proportion of Physicians Ages 65 & Older.

Sources: Left Panel — Authors' Calculation using the LFS, **Right Panel** — Canadian Institute for Health Information.

Figure 3.6: 3-year Moving Average Weekly Hours of Work by Gender and Being Parents.

Source: Authors' Calculation using the LFS.

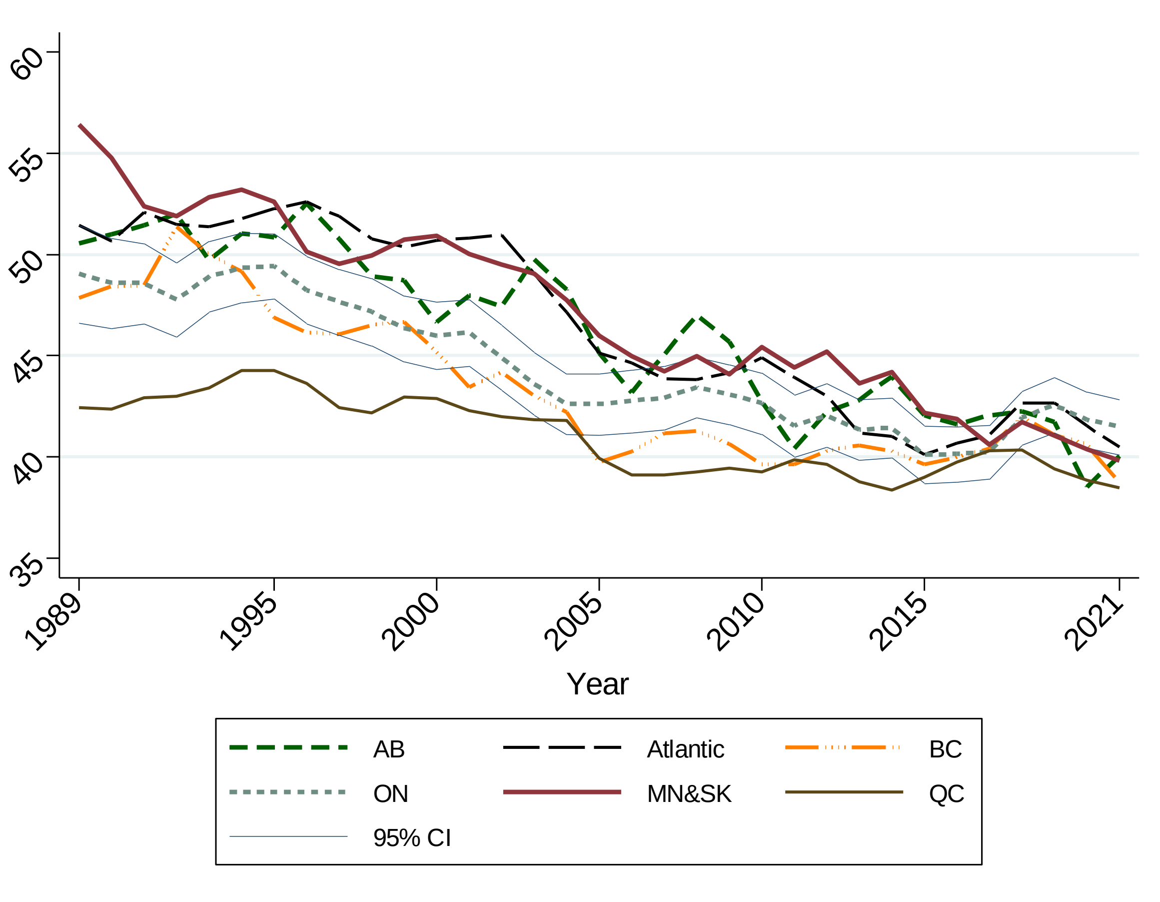


Figure 3.7: 3-year Moving Average Weekly Hours of Work by Region.

Source: Authors' Calculation using the LFS.

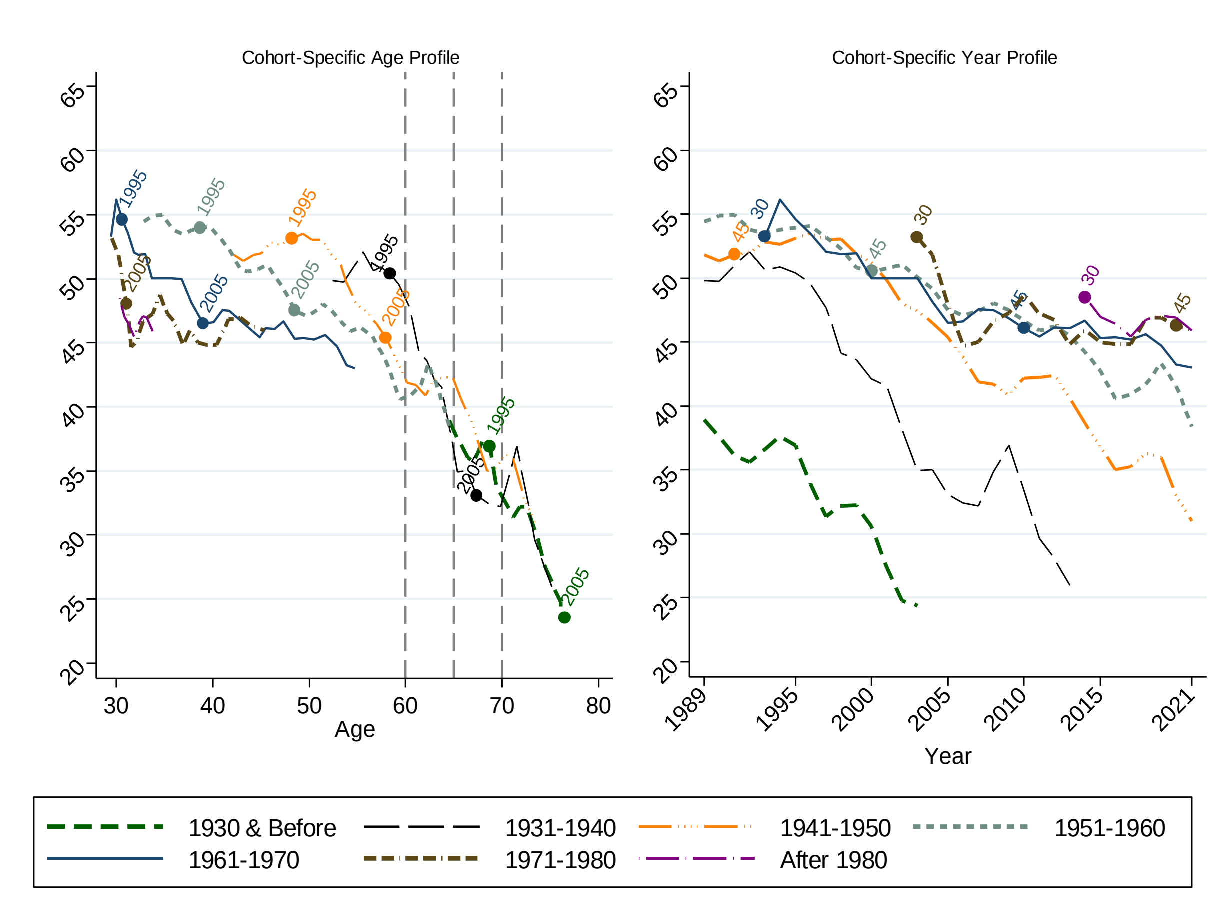


Figure 3.8: 3-year Moving Average Cohort-specific Age and Year Profile of Male Physicians.

Source: Authors' Calculation using the LFS.

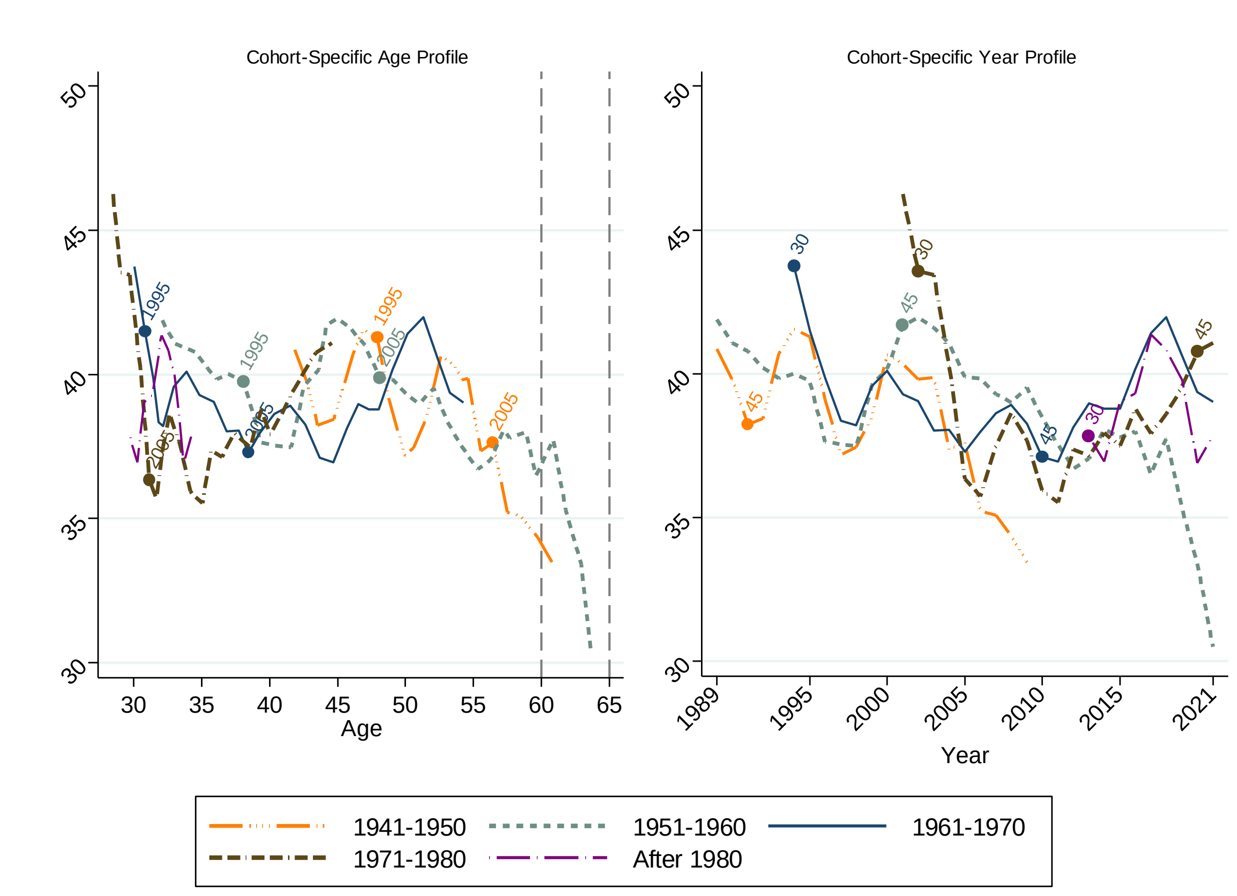


Figure 3.9: 3-year Moving Average Cohort-specific Age and Year Profile of Female Physicians.

Source: Authors' Calculation using the LFS.

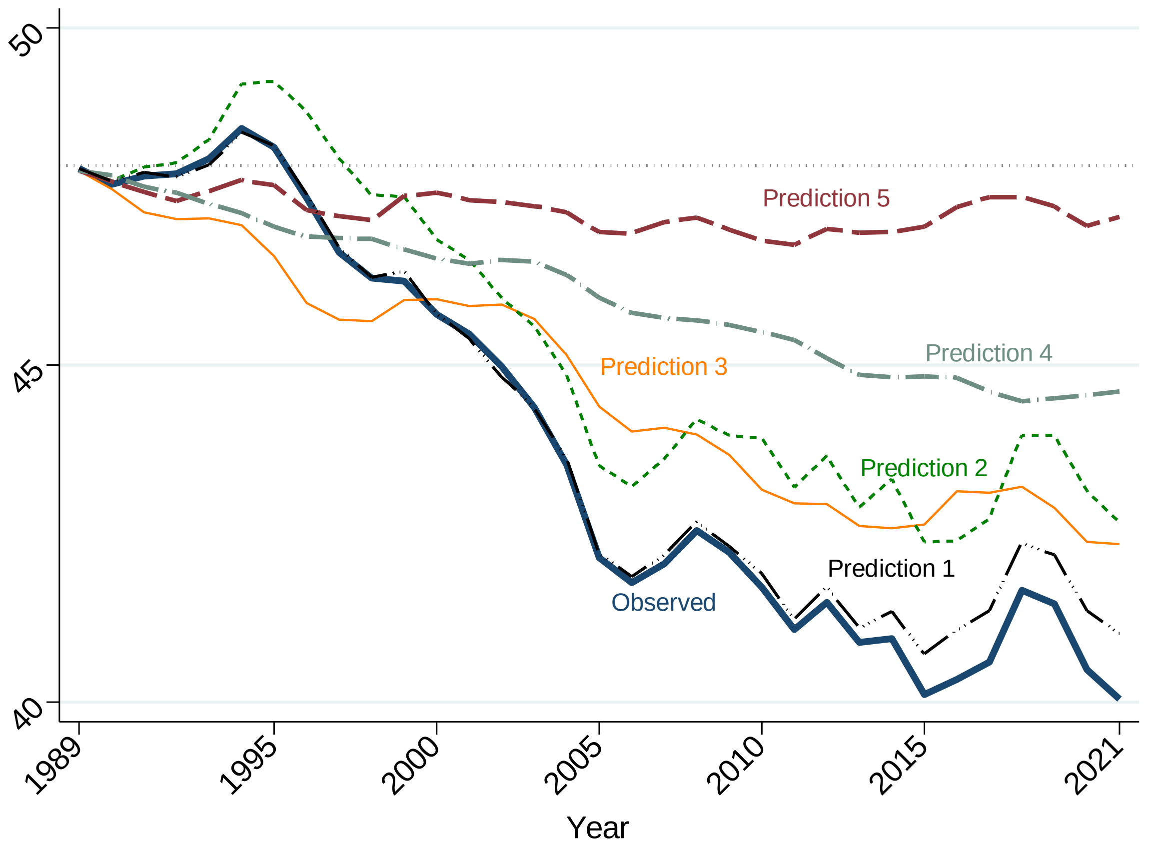


Figure 3.10: Counterfactuals Using the Following Constraints

Prediction 1 holds the proportions of physicians in each of 5 age groups at their 1989 levels.

Prediction 2 holds the proportions of male and female physicians in each of five age groups at the 1989 level, but allows the hours in each group to vary.

Prediction 3 holds average hours of male, but not female, physicians unchanged at the 1989 level, by age group but the share in each age group is allowed to vary.

Prediction 4 holds average hours of both male and female physicians unchanged at the 1989 level, by age group.

Prediction 5 holds the proportions of male, female, and hours of male physicians unchanged at the 1989 level, by age group – that is, only female physicians' hours vary.

Source: Authors' Calculation using the LFS.

Table 3.1: MSEs Using Different Methods for Age Group 65-74

Methods	MSEs
GMM-weighted	1.19
IPF-weighted	2.16
LFS-weighted	3.44
Unweighted	3.14

Table 3.2: Results from OLS Regression Analysis Subsamples Using the Time Intervals

Hours of Work	1987-2021	1987-2000	2001-2021	2001-2010	2011-2021
Female	-4.99***	-4.92***	-4.61***	-4.91***	-4.49***
	(0.58)	(1.03)	(0.71)	(1.08)	(0.93)
40-49	-1.64***	-2.09***	-1.18**	-1.25	-1.23
	(0.44)	(0.69)	(0.57)	(0.79)	(0.81)
50-59	-2.59***	-2.60***	-2.22***	-1.81**	-2.61***
	(0.49)	(0.82)	(0.60)	(0.85)	(0.85)
60-64	-5.57***	-6.22***	-4.99***	-4.91***	-5.09***
	(0.67)	(1.27)	(0.79)	(1.20)	(1.04)
65 & Above	-15.71***	-20.83***	-13.65***	-13.69***	-13.65***
	(0.68)	(1.22)	(0.82)	(1.25)	(1.09)
Child<=5	0.44	1.26	-0.19	1.18	-1.54*
	(0.48)	(0.79)	(0.60)	(0.86)	(0.83)
Child>5	1.82***	2.45***	1.39***	2.26***	0.55
	(0.40)	(0.70)	(0.48)	(0.69)	(0.66)
Hospital	-0.29	-0.99*	0.47	0.86	0.25
	(0.32)	(0.58)	(0.39)	(0.59)	(0.51)
Single/Div/Sep	-0.61*	-0.55	-0.82**	-0.96	-0.68
	(0.35)	(0.73)	(0.40)	(0.63)	(0.52)
40-49 # Female	1.62**	-0.06	1.74*	1.20	2.52**
	(0.75)	(1.39)	(0.90)	(1.32)	(1.23)
50-59 # Female	0.68	-1.91	0.64	0.63	0.78
	(0.83)	(2.02)	(0.96)	(1.48)	(1.26)
60-64 # Female	-1.08	-1.16	-1.47	-3.55	-1.06
	(1.41)	(3.81)	(1.54)	(2.35)	(1.92)
65 & Above # Female	1.47	7.69	-0.59	-1.92	-0.44
	(1.49)	(4.73)	(1.58)	(2.83)	(1.93)
Child<=5 # Female	-9.43***	-12.36***	-8.08***	-9.93***	-6.58***
	(0.77)	(1.37)	(0.93)	(1.42)	(1.23)
Child>5 # Female	-5.61***	-8.43***	-4.58***	-5.61***	-3.68***
	(0.70)	(1.45)	(0.80)	(1.21)	(1.06)
Atlantic	0.86	1.60	0.28	0.11	0.27
	(0.55)	(1.03)	(0.65)	(0.98)	(0.87)
BC	-2.00***	-2.30**	-2.00***	-3.99***	-0.51
	(0.58)	(1.08)	(0.67)	(1.04)	(0.88)
ON	-0.52	-1.57*	-0.20	-1.74*	0.93
	(0.51)	(0.95)	(0.59)	(0.92)	(0.78)
MN & SK	1.21**	2.22**	0.48	0.77	0.17
	(0.58)	(1.10)	(0.67)	(1.03)	(0.87)
QC	-3.24***	-5.85***	-2.10***	-3.92***	-0.73
	(0.53)	(0.96)	(0.63)	(0.97)	(0.83)
Intercept	57.07***	58.55***	53.71***	54.97***	49.53***
	(1.17)	(1.44)	(1.06)	(1.36)	(1.21)
N	95683	31661	64022	27617	36405

*** p<.01", "** p<.05", "* p<.1. Standards errors are in the parentheses.

Note: Survey year and month coefficients are omitted to save space.

Table 3.3: Results from Probit Model for Absences and Negative Binomial for Length of Absences for Full Period

Average Marginal Effect (AME)	Likelihood of Absences	Length of Absences
Female	0.03***	1.44***
	(0.00)	(0.20)
40-49	-0.00	0.03
	(0.00)	(0.21)
50-59	0.01**	0.75**
	(0.00)	(0.32)
60-64	0.02***	0.20
	(0.01)	(0.32)
65 & Above	0.02***	1.37***
	(0.01)	(0.49)
Child<=5	0.02***	1.02***
	(0.00)	(0.24)
Child>5	-0.01*	0.09
	(0.00)	(0.22)
Hospital	0.00	0.33*
	(0.00)	(0.19)
Single/Div/Sep	0.00	-0.13
	(0.00)	(0.20)
Atlantic	-0.01	-0.22
	(0.01)	(0.37)
BC	-0.00	-0.29
	(0.01)	(0.32)
ON	-0.01*	-0.36
	(0.01)	(0.30)
MN & SK	-0.01**	-0.19
	(0.01)	(0.34)
QC	0.00	-0.22
	(0.01)	(0.31)
N	95683	8972

*** p<.01", "** p<.05", "* p<.1. Standards errors are in the parentheses.

Note: Survey year and month coefficients are omitted to save space.

Appendix 3.1: Details of NOCs, NAICS, and Birth Cohorts Applied in the Analysis

NOCs: 3111, specialist physicians; and 3112, general practitioners and family physicians.

NAICS: 2211, offices of physicians; 6214,out-patient care centers; 6215, medical and diagnostic laboratories; 6216, home health care services; 6219, other ambulatory health care services; 6220, hospitals; 6230, nursing and residential care facilities; 9112, other federal government public administration; 9120, provincial and territorial public administration; 6113, universities; 5417, scientific research and development services; 5241, insurance carriers; and 3254, pharmaceutical and medicine manufacturing.

Birth Cohorts: Cohort 1, birth year 1930 or before; Cohort 2, birth year between 1931-1940; Cohort 3, birth year between 1941-1950; Cohort 4, birth year between 1951-1960; Cohort 5, birth year between 1961-1970; Cohort 6, birth year between 1971-1980; and Cohort 7, birth year after 1980.

Appendix 3.2: Proof that 1-step and 2-step Implementations of H&I's Approach Are Equivalent for Simple Means in Certain Circumstances

Let $G_\mu = \frac{\partial}{\partial \mu_y} G(\mu_y, \lambda), G_{\lambda_j} = \frac{\partial}{\partial \lambda_j} G(\mu_y, \lambda), F_{i,\lambda_j} = \frac{\partial}{\partial \lambda_j} F_i(\lambda)$, $\theta = \begin{bmatrix} \mu_y \\ \lambda_1 \\ \vdots \\ \lambda_I \end{bmatrix}$ and

$\Psi(\theta) = \begin{bmatrix} G(\mu_y, \lambda) \\ F_1(\lambda) \\ \vdots \\ F_I(\lambda) \end{bmatrix}$, then the Jacobian matrix for the 1-step approach is

$$J(\theta) \coloneqq \frac{\partial \Psi(\theta)}{\partial \theta} = \begin{bmatrix} G_\mu & G_{\lambda_1} & \dots & G_{\lambda_I} \\ 0 & F_{1,\lambda_1} & \dots & F_{1,\lambda_I} \\ \vdots & \vdots & \ddots & \vdots \\ 0 & F_{I,\lambda_1} & \dots & F_{I,\lambda_I} \end{bmatrix}.$$

Focusing on the Gauss-Newton algorithm, at any given iteration s (including the initial step with common initial values) updates the estimates as follows:

$$\theta^{(s+1)} = \theta^{(s)} - \left(J(\theta^{(s)})^T J(\theta^{(s)}) \right)^{-1} J(\theta^{(s)})^T \Psi(\theta^{(s)}).$$

When we follow a 2-step procedure and estimate only λ, the first row and first column of the Jacobian matrix, $J(\theta)$, are eliminated. We write the expression as:

$$\Lambda(\lambda) \coloneqq \begin{bmatrix} F_{1,\lambda_1} & \cdots & F_{1,\lambda_I} \\ \vdots & \ddots & \vdots \\ F_{I,\lambda_1} & \cdots & F_{I,\lambda_I} \end{bmatrix}.$$

Note that $F_{i,\lambda_j} = F_{j,\lambda_i}, \forall i, j$. Therefore, $\Lambda(\lambda)$ is symmetric. The Gauss-Newton algorithm admits the following form:

$$\tilde{\lambda}^{(s+1)} = \tilde{\lambda}^{(s)} - \left(\Lambda\left(\tilde{\lambda}^{(s)}\right)^T \Lambda\left(\tilde{\lambda}^{(s)}\right)\right)^{-1} \Lambda\left(\tilde{\lambda}^{(s)}\right)^T \mathrm{F}\left(\tilde{\lambda}^{(s)}\right).$$

Our goal is to show both 1-step and 2-step yield the same estimates, given the same initial values for λ.

First compute $\left(J\left(\theta^{(s)}\right)^T J\left(\theta^{(s)}\right)\right)^{-1}$. Assuming $C(M)$ to be a cofactor matrix of any given full-ranked square matrix M, we can write:

$$\left(J\left(\theta^{(s)}\right)^T J\left(\theta^{(s)}\right)\right)^{-1} = \frac{1}{(\det[J(\theta^{(s)})])^2}\left(C\left(J\left(\theta^{(s)}\right)^T\right) C\left(J\left(\theta^{(s)}\right)\right)\right)^T,$$

where $\det\left[J\left(\theta^{(s)}\right)\right] = G_\mu \det[\Lambda(\lambda)]$. These quantities are evaluated at iteration s (i.e., at values $\theta^{(s)}$).

Denoting $\left|\Lambda_{i,j}\right|$ as the determinant of Λ after eliminating i-th row and j-th column, we can write cofactor of $J(\theta)$ as follows. For instance, a cofactor of $(2,1)$-th element of $J(\theta)$ is:

$$(-1)^{2+1} \det\begin{bmatrix} G_{\lambda_1} & G_{\lambda_2} & \dots & G_{\lambda_I} \\ F_{2,\lambda_1} & F_{2,\lambda_2} & \dots & F_{2,\lambda_I} \\ \vdots & \vdots & \ddots & \vdots \\ F_{I,\lambda_1} & F_{I,\lambda_2} & \dots & F_{I,\lambda_I} \end{bmatrix} = (-1)^{2+1} \sum_{i=1}^{I} (-1)^{1+i}\, G_{\lambda_i} \left|\Lambda_{1,i}\right|.$$

Then,

$$C\left(J\left(\theta^{(s)}\right)\right) = \begin{bmatrix} \det(\Lambda) & 0 & \dots & 0 \\ (-1)^{2+1} \sum_{i=1}^{I} (-1)^{1+i}\, G_{\lambda_i} \left|\Lambda_{1,i}\right| & (-1)^{2+2} G_\mu \left|\Lambda_{1,1}\right| & \dots & (-1)^{2+I+1} G_\mu \left|\Lambda_{1,I}\right| \\ \vdots & \vdots & \ddots & \vdots \\ (-1)^{I+1+1} \sum_{i=1}^{I} (-1)^{1+i}\, G_{\lambda_i} \left|\Lambda_{I,i}\right| & (-1)^{I+1+2} G_\mu \left|\Lambda_{I,1}\right| & \dots & (-1)^{I+1+I+1} G_\mu \left|\Lambda_{I,I}\right| \end{bmatrix}$$

And the transpose,

$$C\left(J\left(\theta^{(s)}\right)^T\right) = \begin{bmatrix} \det(\Lambda) & (-1)^{1+2}\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{1,i}^T\right| & \dots & (-1)^{1+I+1}\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{I,i}^T\right| \\ 0 & (-1)^{2+2}G_\mu\left|\Lambda_{1,1}^{\mathrm{T}}\right| & \dots & (-1)^{2+I+1}G_\mu\left|\Lambda_{1,I}^T\right| \\ \vdots & \vdots & \ddots & \vdots \\ 0 & (-1)^{I+1+2}G_\mu\left|\Lambda_{I,1}^T\right| & \dots & (-1)^{I+1+I+1}G_\mu\left|\Lambda_{I,I}^{\mathrm{T}}\right| \end{bmatrix}.$$

As Λ is symmetric, $\left|\Lambda_{i,j}^T\right| = \left|\Lambda_{i,j}\right|, \forall i,j$. Hence,

$$C\left(J\left(\theta^{(s)}\right)^T\right) = \begin{bmatrix} \det(\Lambda) & (-1)^{1+2}\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{1,i}\right| & \dots & (-1)^{1+I+1}\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{I,i}\right| \\ 0 & (-1)^{2+2}G_\mu\left|\Lambda_{1,1}\right| & \dots & (-1)^{2+I+1}G_\mu\left|\Lambda_{1,I}\right| \\ \vdots & \vdots & \ddots & \vdots \\ 0 & (-1)^{I+1+2}G_\mu\left|\Lambda_{I,1}\right| & \dots & (-1)^{I+1+I+1}G_\mu\left|\Lambda_{I,I}\right| \end{bmatrix}$$

Therefore,

$$C\left(J\left(\theta^{(s)}\right)^T\right)C\left(J\left(\theta^{(s)}\right)\right) =$$

$$\begin{bmatrix} \det(\Lambda)^2 + \sum_{j=1}^{I}\left(\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right|\right)^2 & G_\mu\sum_{j=1}^{I}(-1)^{2(j+1)+1+2}\left|\Lambda_{j,1}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & \dots & G_\mu\sum_{j=1}^{I}(-1)^{2(j+1)+1+(I+1)}\left|\Lambda_{j,I}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| \\ G_\mu\sum_{j=1}^{I}(-1)^{2(j+1)+2+1}\left|\Lambda_{j,1}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & G_\mu^2\sum_{j=1}^{I}(-1)^{2(j+1)+2+2}\left|\Lambda_{j,1}\right|^2 & \dots & G_\mu^2\sum_{j=1}^{I}(-1)^{2(j+1)+2+(I+1)}\left|\Lambda_{j,1}\right|\left|\Lambda_{j,I}\right| \\ \vdots & \vdots & \ddots & \vdots \\ G_\mu\sum_{j=1}^{I}(-1)^{2(j+1)+(I+1)+1}\left|\Lambda_{j,I}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & G_\mu^2\sum_{j=1}^{I}(-1)^{2(j+1)+2+(I+1)}\left|\Lambda_{j,I}\right|\left|\Lambda_{j,1}\right| & \dots & G_\mu^2\sum_{j=1}^{I}(-1)^{2(j+1)+(I+1)+(I+1)}\left|\Lambda_{j,I}\right|^2 \end{bmatrix}$$

$$= \begin{bmatrix} \det(\Lambda)^2 + \sum_{j=1}^{I}\left(\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right|\right)^2 & G_\mu\sum_{j=1}^{I}(-1)^{1+2}\left|\Lambda_{j,1}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & \dots & G_\mu\sum_{j=1}^{I}(-1)^{1+(I+1)}\left|\Lambda_{j,I}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| \\ G_\mu\sum_{j=1}^{I}(-1)^{2+1}\left|\Lambda_{j,1}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & G_\mu^2\sum_{j=1}^{I}(-1)^{2+2}\left|\Lambda_{j,1}\right|^2 & \dots & G_\mu^2\sum_{j=1}^{I}(-1)^{2+(I+1)}\left|\Lambda_{j,1}\right|\left|\Lambda_{j,I}\right| \\ \vdots & \vdots & \ddots & \vdots \\ G_\mu\sum_{j=1}^{I}(-1)^{(I+1)+1}\left|\Lambda_{j,I}\right|\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}\left|\Lambda_{j,i}\right| & G_\mu^2\sum_{j=1}^{I}(-1)^{2+(I+1)}\left|\Lambda_{j,I}\right|\left|\Lambda_{j,1}\right| & \dots & G_\mu^2\sum_{j=1}^{I}(-1)^{(I+1)+(I+1)}\left|\Lambda_{j,I}\right|^2 \end{bmatrix}$$

This is a $(I+1)\times(I+1)$ matrix. Each element in the r-th row and c-th column (r,c) can be expressed as:

$$a_{(r,c)} =$$

$$\begin{cases} G_\mu^2 \sum_{j=1}^{I} (-1)^{r+c} |\Lambda_{j,r-1}|\,|\Lambda_{j,c-1}|, & \text{if } r>1 \text{ and } c>1 \\ G_\mu \sum_{j=1}^{I} (-1)^{r+c} |\Lambda_{j,\max\{r,c\}}| \sum_{i=1}^{I} (-1)^{1+i}\, G_{\lambda_i} |\Lambda_{j,i}|, & \text{if } \{r=1 \text{ and } c>1\} \text{ or } \{r>1 \text{ and } c=1\} \end{cases}$$

Now,

$$\left(J(\theta^{(s)})^T J(\theta^{(s)})\right)^{-1} J(\theta^{(s)})^T = \frac{1}{(\det[J(\theta^{(s)})])^2}\left(C\left(J(\theta^{(s)})^T\right) C\left(J(\theta^{(s)})\right)\right)^T J(\theta^{(s)})^T$$
$$= \frac{1}{(\det[J(\theta^{(s)})])^2}\left[b_{(r,c)}\right]_{r,c=1,\dots,I+1} \qquad (J1)$$

where

$$b_{(1,1)} = G_\mu \det(\Lambda)^2 + G_\mu \sum_{j=1}^{I}\left(\sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}|\Lambda_{j,i}|\right)^2$$
$$+ G_\mu \sum_{k=1}^{I} G_{\lambda_k} \sum_{j=1}^{I}(-1)^{1+(k+1)}|\Lambda_{j,k}| \sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}|\Lambda_{j,i}|$$

$$b_{(1,2)} = G_\mu \sum_{k=1}^{I} F_{1,\lambda_k} \sum_{j=1}^{I}(-1)^{1+(k+1)}|\Lambda_{j,k}| \sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}|\Lambda_{j,i}|$$

$$b_{(1,I+1)} = G_\mu \sum_{k=1}^{I} F_{I,\lambda_k} \sum_{j=1}^{I}(-1)^{1+(k+1)}|\Lambda_{j,k}| \sum_{i=1}^{I}(-1)^{1+i}\, G_{\lambda_i}|\Lambda_{j,i}|$$

$$b_{(2,1)} = G_\mu^2 \sum_{j=1}^{I} (-1)^{2+1} |\Lambda_{j,1}| \sum_{i=1}^{I} (-1)^{1+i}\, G_{\lambda_i} |\Lambda_{j,i}| + G_\mu^2 \sum_{k=1}^{I} G_{\lambda_k} \sum_{i=1}^{I} (-1)^{2+(k+1)}\, |\Lambda_{j,k}||\Lambda_{j,1}|$$

$$=^{k\to i}\; G_\mu^2 \sum_{j=1}^{I} \sum_{i=1}^{I} |\Lambda_{j,1}||\Lambda_{j,i}| G_{\lambda_i} \underbrace{\left[(-1)^{4+i} + (-1)^{3+i}\right]}_{=0} = 0$$

The last expression is analogous to all $b_{(r,1)}\ with\, r \geq 2$, which are all zeroes. Further,

$$b_{(2,2)} = G_\mu^2 \sum_{k=1}^{I} F_{1,\lambda_k} \sum_{j=1}^{I} (-1)^{2+(k+1)} |\Lambda_{j,1}|\, |\Lambda_{j,k}|.$$

In general, for any pair (r, c) with $r > 1, c > 1$

$$b_{(r,c)} = G_\mu^2 \sum_{k=1}^{I} F_{c-1,\lambda_k} \sum_{j=1}^{I} (-1)^{r+(k+1)} |\Lambda_{j,r-1}|\, |\Lambda_{j,k}|.$$

Recall equation (*J1*) that gives the expression $\frac{1}{\left(\det[J(\theta^{(s)})]\right)^2}\left[b_{(r,c)}\right]_{r,c=1,\ldots,I+1}$. By multiplying each element of matrix $\left[b_{(r,c)}\right]_{r,c=1,\ldots,I+1}$ by $\frac{1}{\left(\det[J(\theta^{(s)})]\right)^2}\ or\ \frac{1}{G_\mu^2 \det(\Lambda)^2}$, we find $\theta^{(s+1)}$depends neither on any derivatives of $G(\cdot)$ nor on $\mu_y^{(s)}$.

Under the same analogy, it is easy to show that $\lambda^{(s+1)}$ from vector $\theta^{(s+1)}$ is exactly the same as it from vector $\tilde{\lambda}^{(s+1)}$ since $\frac{1}{G_\mu^2 \det(\Lambda)^2}\left[b_{(r,c)}\right]_{r,c=2,\ldots,I+1}$ is the same in both cases. These imply that, given the same initial values for λ, the values of estimates $\lambda^{(s)}$ are the same in each iteration s of the Gauss-Newton algorithm regardless of inclusion of the moment for μ_y.

Note that the exact numerical equivalence does not apply to all solution algorithms. For example, if the Newton-Rapson were employed instead, the 1- and 2-step approaches would provide very similar results in finite samples, but they would not be numerically identical.

Appendix 3.3: Deviation of the Sample Data from Population Data and Monte Carlo Simulation

Appendix 3.3.1: Comparisons of the LFS to CIHI Data

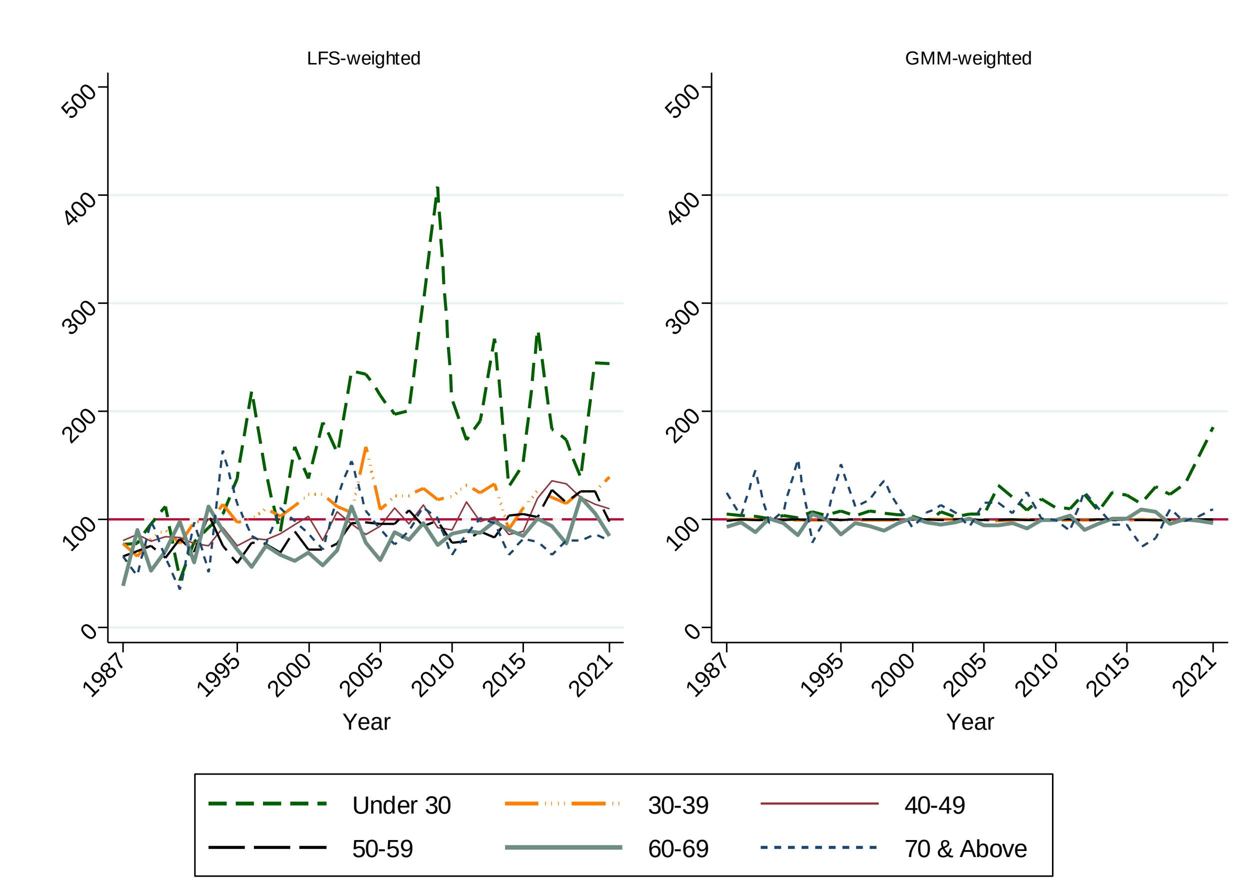


Figure A3.1: Comparison of Proportion of Physicians Sampled by Age, Using LFS and GMM Weights.

Source: Authors' Calculation using the LFS and CIHI data.

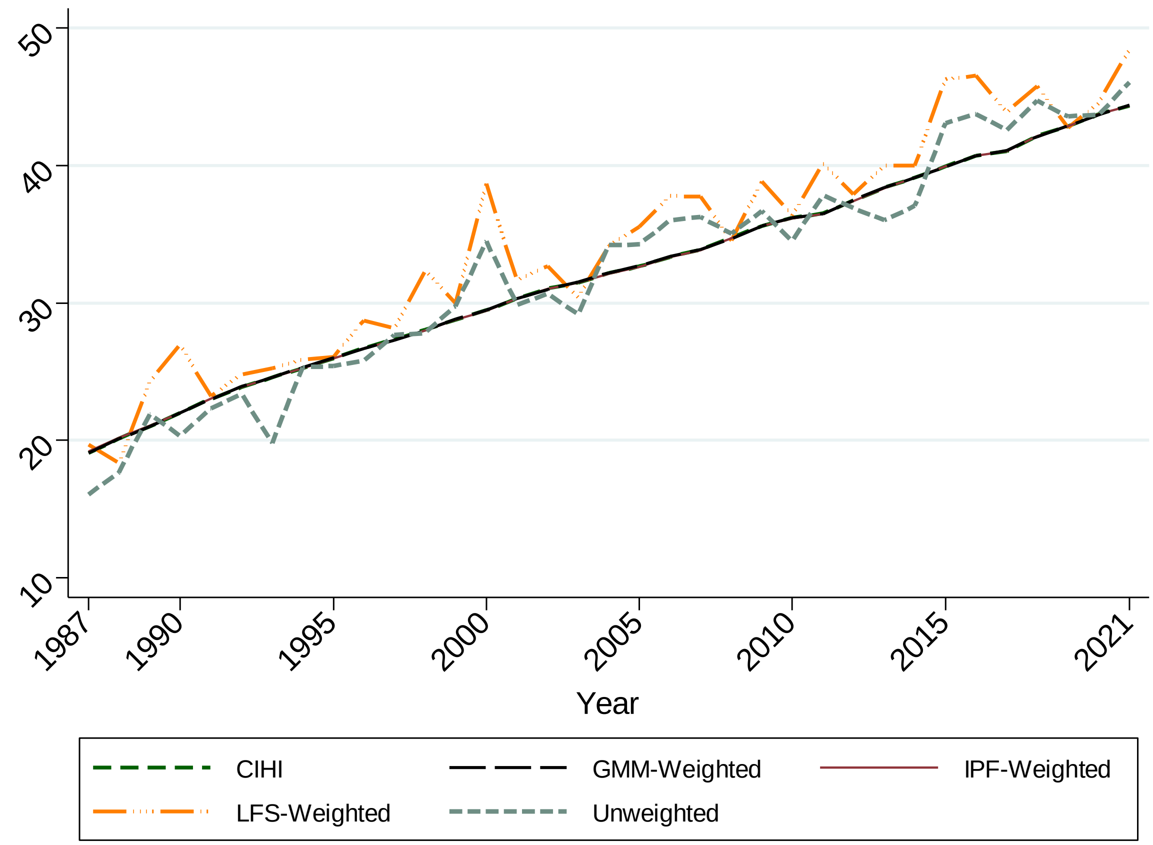


Figure A3.2: Comparison of the Proportion of Female Physicians, Using LFS, GMM and IPF Weights, and No Weights.

Source: Authors' Calculation using the LFS and CIHI data.

Appendix 3.3.2: Simulation Setting

In MCS of coefficient estimates, from pooled data between the years 1987-2020, we construct moments and obtain count totals of males and females in Canada, by each age group in each region, by gender in each region, and by the average age of physicians in Canada to impose restriction in the GMM & IPF. These moments are constructed on the same variables that we apply to reweighting in our samples in question using the CIHI data. Our population size (around 93000) coincidentally reaches almost the same size as the present physician population in Canada. These pooled data enable us to maintain similar real-world correlations among variables as in our analysis samples, further strengthening the conclusion. We draw samples of sizes 850, 1700, and 3500 selecting on gender and geography. As presented in Table A3.1, our population has 67% male and 33% female, but the selected sample has 50% male and 50% female physicians, clearly oversampling females. Further, we undersample in Ontario, Quebec, and Atlantic regions and oversample in the other regions.

Table A3.1: The Population and Sample Proportions of Male and Female Physicians by Region

				Population Proportion			
	AB	Atlantic	BC	ON	MN & SK	QC	Total
Male (%)	66.21	68.23	67.61	66.53	70.19	63.18	66.81
Female (%)	33.79	31.77	32.39	33.47	29.81	36.84	33.19
Region (%)	8.14	17.40	11.37	30.45	14.07	18.57	100
				Sample Proportion			
Male (%)	50	50	50	50	50	50	50
Female (%)	50	50	50	50	50	50	50
Region (%)	11.77	14.66	17.66	23.54	17.65	14.72	100

Appendix 3.3.3: More Simulation Results

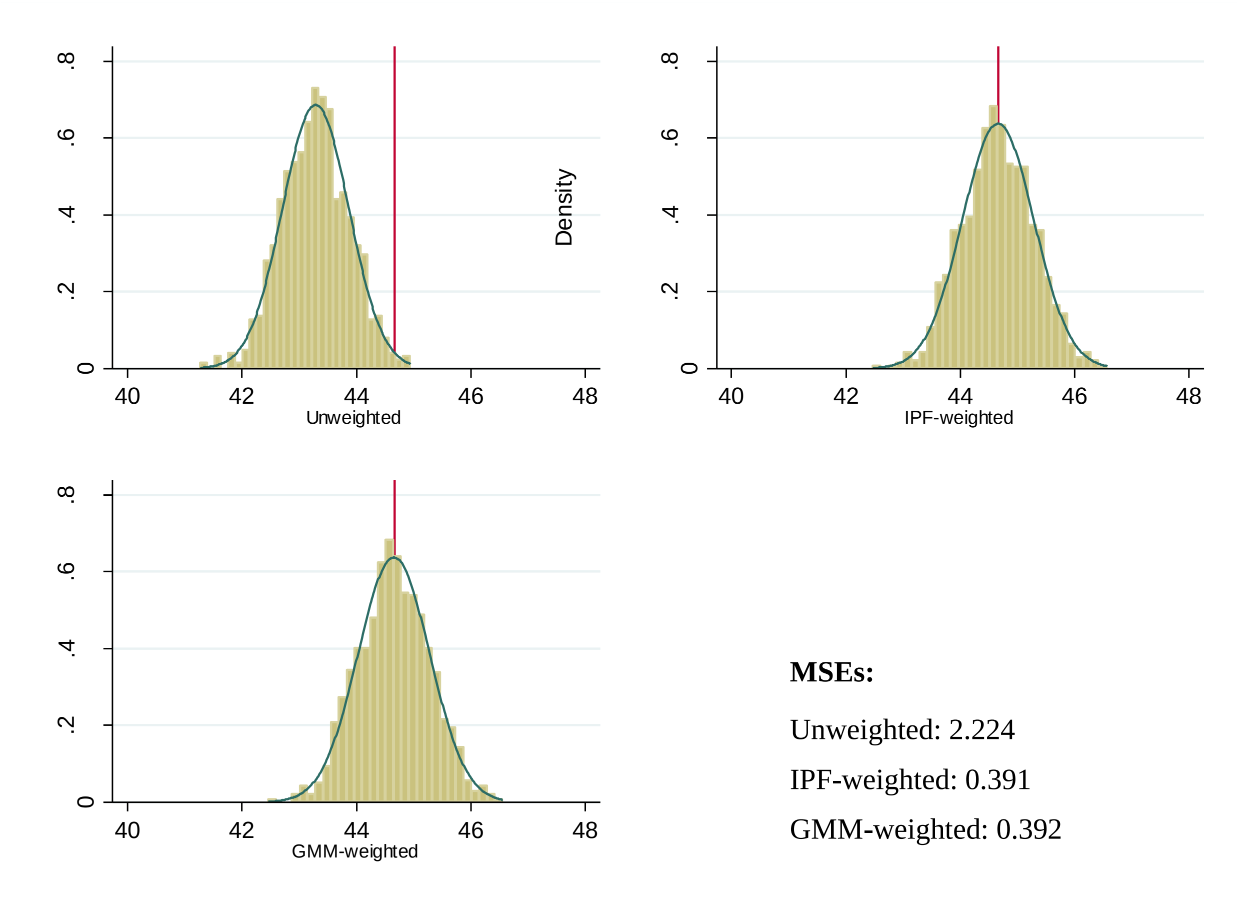


Figure A3.3: Distribution of Weighted Average Hours in MCS Using the Sample Size of 1700 with Same Information in the GMM & IPF.

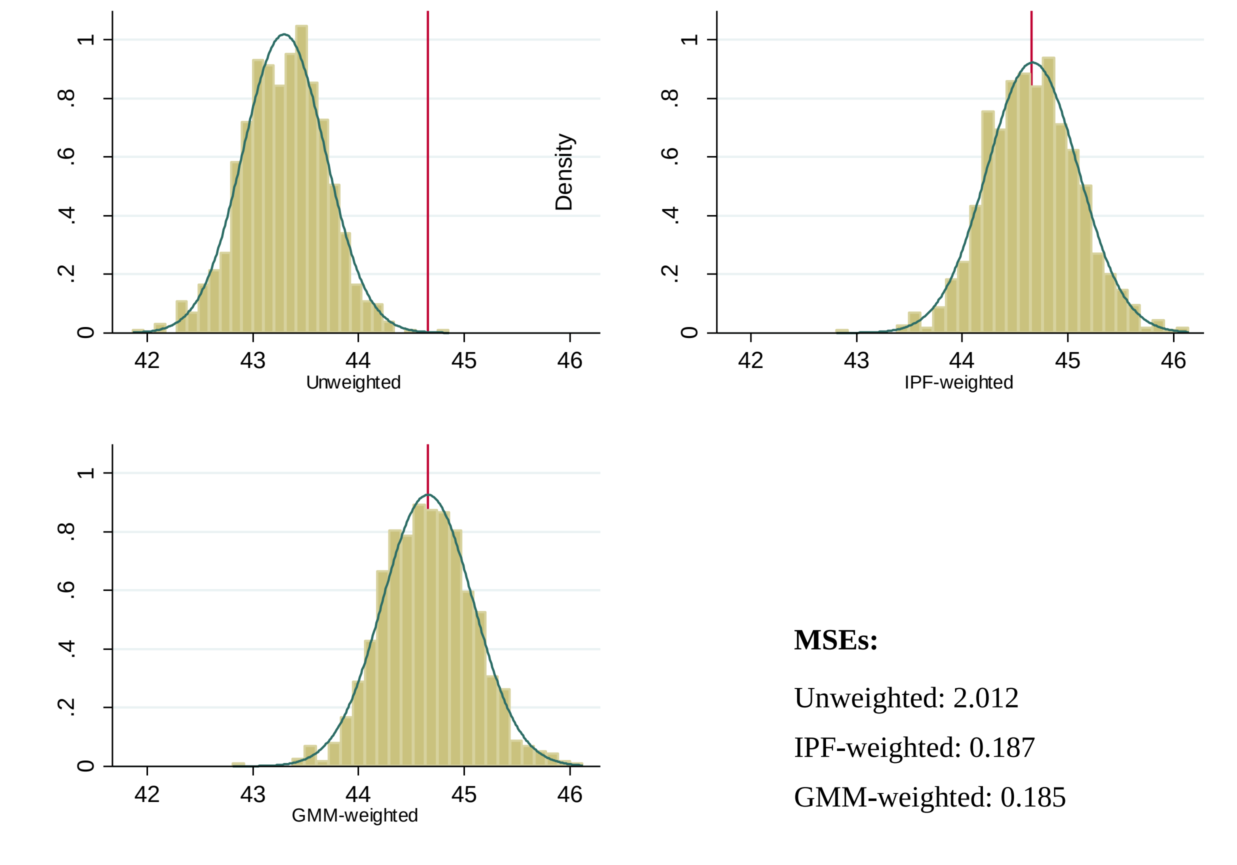


Figure A3.4: Distribution of Weighted Average Hours in MCS Using the Sample Size of 3500 with Mixed Continuous and Discrete Constraints for the GMM.

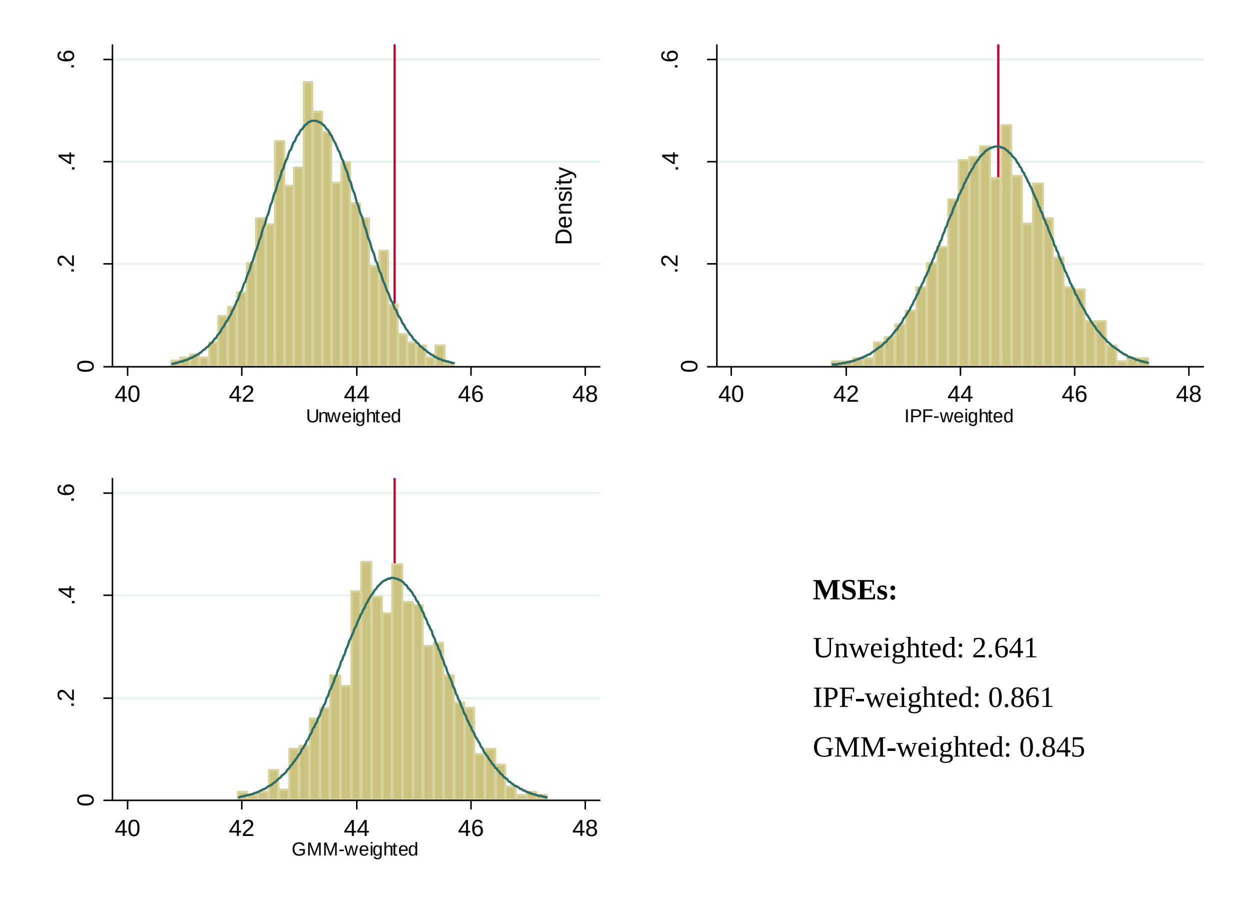


Figure A3.5: Distribution of Weighted Average Hours in MCS Using the Sample Size of 850 with Mixed Continuous and Discrete Constraints for the GMM.

Appendix 3.3.4: Alternative Specifications

We conduct two additional sets of Monte Carlo simulations using alternative specifications using a sample size of 1700 with same sampling technique as in the main simulations. In the first alternative, we reduce the number of constraints from GMM, i.e., we use the mean of the continuous variable age with four of the five age categories in each region, ignoring the fifth category, and incorporating the proportions of females in Canada and five of the six regions. In Table A3.2, we present MSEs of this alternative. The table shows that the MSEs from 1000 repetitions is still minimum in the GMM (0.342) and highest in the unweighted estimates (2.194), though the difference between IPF and GMM is trivial. In the second alternative, we increase the number of constraints in IPF with additional information on the count totals of each age group in Canada. In this alternative, we keep the number of constraints unchanged for the GMM, i.e., we use the mean of the continuous variable age with four of the five age categories in each region and other category (below 30) for four regions, the proportion of females in Canada and five of the six regions. The results show that GMM still has the minimum MSEs, and the gap remains the same as in the first alternative. Therefore, we do not present the results from the second alternative here to save the space.

Table A3.2**:** MSEs of Different Methods in Monte Carlo Simulation in the Firs Alternative

Methods	MSEs
GMM-weighted	0.342
IPF-weighted	0.346
Unweighted	2.194

Appendix 3.4: Ancillary Results, and Calculation of Counterfactuals

Appendix 3.4.1: Results

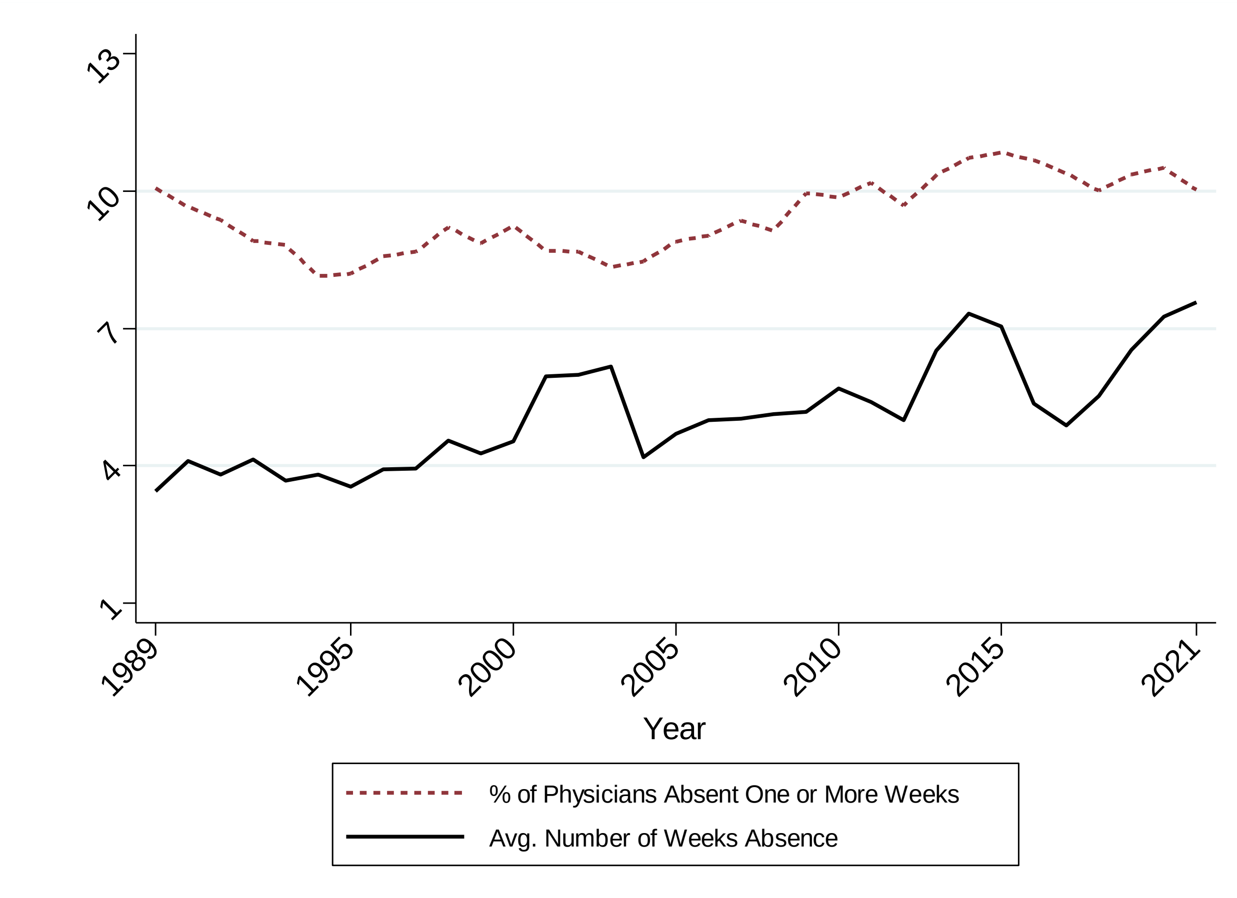


Figure A3.6: 3-Year Moving Average Percent Physicians Absent and Length of Absences from Work.

Source: Authors' Calculation using the LFS.

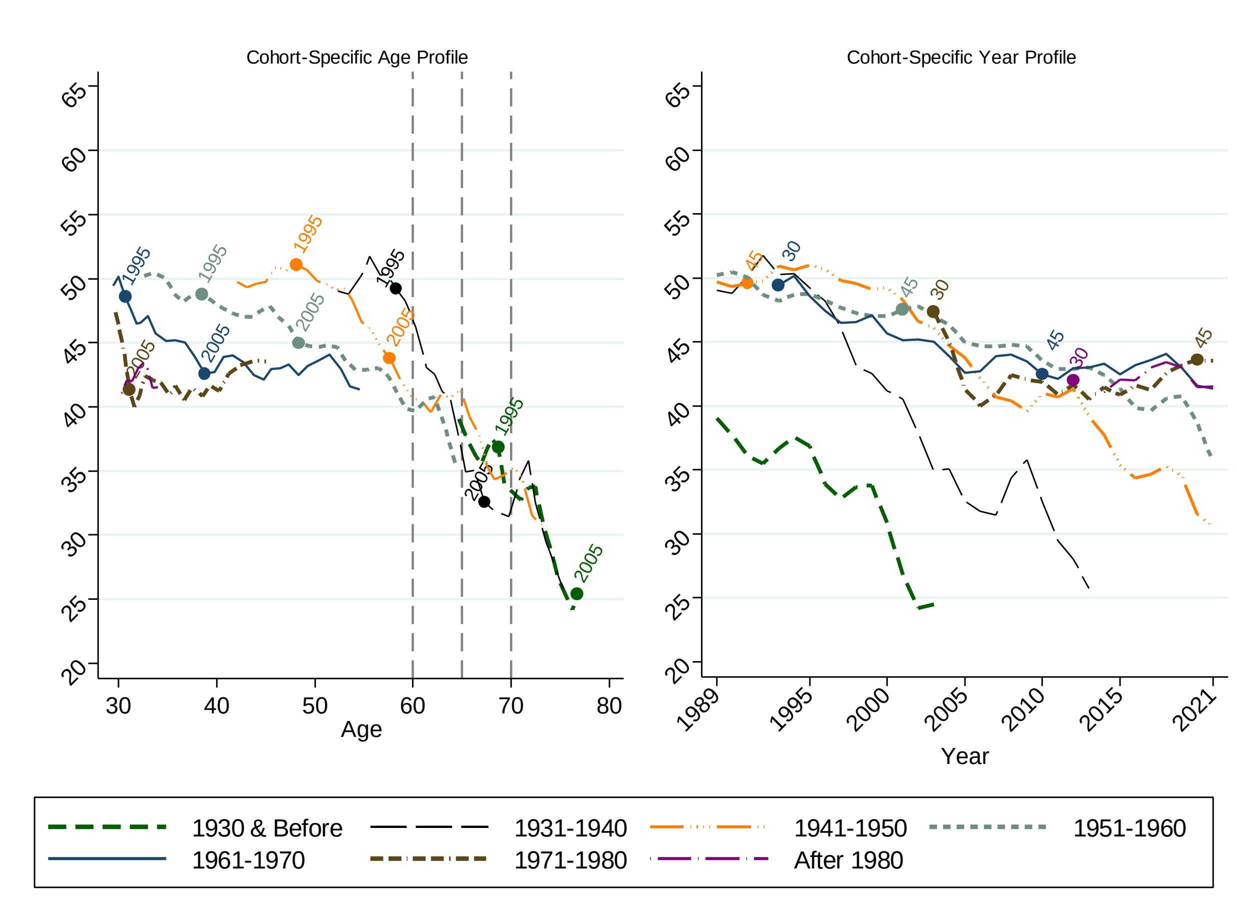


Figure A3.7: 3-Year Moving Average Cohort-specific Age and Year Profile of All Physicians.

Source: Authors' Calculation using the LFS.

Appendix 3.4.2: Calculation of Counterfactuals

Counterfactuals are calculated using 3-year moving average hours of work data, in which two lags and current periods are included. Due to 3-year moving average, we lose data for the first two years, which is why we impose restrictions on 1989 instead of 1987. H_{c1}- H_{c5} show how we calculate counterfactuals 1-5.

H_{c1} holds the proportions of physicians in each of 5 age groups at their 1989 levels.

$$H_{c1} = \frac{Phys2839_{1989}}{TotPhys_{1989}} * AvgHrs2839_t + \frac{Phys4049_{1989}}{TotPhys_{1989}} * AvgHrs4049_t + \frac{Phys5059_{1989}}{TotPhys_{1989}} * AvgHrs5059_t + \frac{Phys6064_{1989}}{TotPhys_{1989}} * AvgHrs6064_t + \frac{Phys65p_{1989}}{TotPhys_{1989}} * AvgHrs65p_t$$

H_{c2} holds the proportions of male and female physicians in each of five age groups at the 1989 level, but allows hours the hours in each group to vary.

$$H_{c2} = \frac{Male2839_{1989}}{TotPhys_{1989}} * MaleAvgHrs2839_t + \frac{Fem2839_{1989}}{TotPhys_{1989}} * FemAvgHrs2839_t + \frac{Male4049_{1989}}{TotPhys_{1989}} * MaleAvgHrs4049_t + \frac{Fem4049_{1989}}{TotPhys_{1989}} * FemAvgHrs4049_t + \frac{Male5059_{1989}}{TotPhys_{1989}} * MaleAvgHrs5059_t + \frac{Fem5059_{1989}}{TotPhys_{1989}} * FemAvgHrs5059_t + \frac{Male6064_{1989}}{TotPhys_{1989}} * MaleAvgHrs6064_t + \frac{Fem6064_{1989}}{TotPhys_{1989}} * FemAvgHrs6064_t + \frac{Male65p_{1989}}{TotPhys_{1989}} * MaleAvgHrs65p_t + \frac{Fem65p_{1989}}{TotPhys_{1989}} * FemAvgHrs65p_t$$

H_{c3} holds average hours of male, but not female, physicians unchanged at the 1989 level, by age group but the share in each age group is allowed to vary.

$$H_{c3} = \frac{Male2839_t}{TotPhys_t} * MaleAvgHrs2839_{1989} + \frac{Fem2839_t}{TotPhys_t} * FemAvgHrs2839_t + \frac{Male4049_t}{TotPhys_t} * MaleAvgHrs4049_{1989} + \frac{Fem4049_t}{TotPhys_t} * FemAvgHrs4049_t + \frac{Male5059_t}{TotPhys_t} * MaleAvgHrs5059_{1989} + \frac{Fem5059_t}{TotPhys_t} * FemAvgHrs5059_t + \frac{Male6064_t}{TotPhys_t} * MaleAvgHrs6064_{1989} + \frac{Fem6064_t}{TotPhys_t} * FemAvgHrs6064_t + \frac{Male65p_t}{TotPhys_t} * MaleAvgHrs65p_{1989} + \frac{Fem65p_t}{TotPhys_t} * FemAvgHrs65p_t$$

H_{c4} holds average hours of both male and female physicians unchanged at the 1989 level, by age group.

$$H_{c4} = \frac{Male2839_t}{TotPhys_t} * MaleAvgHrs2839_{1989} + \frac{Fem2839_t}{TotPhys_t} * FemAvgHrs2839_{1989} + \frac{Male4049_t}{TotPhys_t} * MaleAvgHrs4049_{1989} + \frac{Fem4049_t}{TotPhys_t} * FemAvgHrs4049_{1989} + \frac{Male5059_t}{TotPhys_t} * MaleAvgHrs5059_{1989} + \frac{Fem5059_t}{TotPhys_t} * FemAvgHrs5059_{1989} + \frac{Male6064_t}{TotPhys_t} * MaleAvgHrs6064_{1989} + \frac{Fem6064_t}{TotPhys_t} * FemAvgHrs6064_{1989} + \frac{Male65p_t}{TotPhys_t} * MaleAvgHrs65p_{1989} + \frac{Fem65p_t}{TotPhys_t} * FemAvgHrs65p_{1989}$$

H_{c5} holds the proportions of male, female, and hours of male physicians unchanged at the 1989 level, by age group – that is, only female physicians' hours vary.

$$H_{c5} = \frac{Male2839_{1989}}{TotPhys_{1989}} * MaleAvgHrs2839_{1989} + \frac{Fem2839_{1989}}{TotPhys_{1989}} * FemAvgHrs2839_t + \frac{Male4049_{1989}}{TotPhys_{1989}} * MaleAvgHrs4049_{1989} + \frac{Fem4049_{1989}}{TotPhys_{1989}} * FemAvgHrs4049_t + \frac{Male5059_{1989}}{TotPhys_{1989}} * MaleAvgHrs5059_{1989} + \frac{Fem5059_{1989}}{TotPhys_{1989}} * FemAvgHrs5059_t + \frac{Male6064_{1989}}{TotPhys_{1989}} * MaleAvgHrs6064_{1989} + \frac{Fem6064_{1989}}{TotPhys_{1989}} * FemAvgHrs6064_t + \frac{Male65p_{1989}}{TotPhys_{1989}} * MaleAvgHrs65p_{1989} + \frac{Fem65p_{1989}}{TotPhys_{1989}} * FemAvgHrs65p_t$$

Conclusion

This thesis explores the origins of physician shortages and the declining trends in physician labour supply by analyzing historical data concerning physician work hours and population service demands. Additionally, it investigates physicians' responses to the COVID-19 pandemic, aiming to offer insights into enhancing preparedness for future health crises of similar or lesser magnitudes. The research establishes a connection between the pre-existing shortage and the impact of COVID-19 on future workforce planning. In the first paper, I examine why Canada struggles with physician shortages despite having historically high physician-to-population ratios. The subsequent paper explores the question: Did physicians respond unfavorably to COVID-19, thereby exacerbating the existing shortages? Lastly, the third paper examines the factors contributing to Canada's declining patterns in physician labour supply.

The findings of the first paper show that per capita physician numbers lead to misleading policy formulation because neither average work hours by physicians nor average service demand by the population has been stable over time. Physicians' average work hours declined by about 20%, while demand increased by about 66% during the study period due to population aging. This declining labour supply pattern and the increasing service demand must be considered in planning the future physician workforce. Therefore, the first chapter offers a new perspective not addressed before in offering policy suggestions to maintain an appropriate-sized physician workforce. However, this paper cannot separate direct patient care from the rest of physicians' work, e.g., research and teaching, as the data provide only total weekly hours of work.

The second paper, contrary to media reports and public perceptions, finds physicians worked at the pre-pandemic level during COVID-19 except for the first wave. The reduction happened entirely in the community setting, while in the hospital setting, work hours remained stable. No statistically significant differences were observed across general practitioners/family physicians (GP/FP) and specialists or across genders. Senior physicians aged 60 and above responded more by reducing their hours than their younger counterparts, highlighting the vulnerability of physicians aging during a crisis. The regional variations were also noticeable, with Alberta showing the highest and Quebec the lowest reduction.

Despite the challenges associated with the pandemic, physicians did not escalate their rate or length of absence in any of the waves of COVID-19. Instead, female physicians reduced their probability of being absent from work in the third wave of the pandemic. Furthermore, this research does not find evidence of accelerated retirement among physicians due to COVID-19, though the study could not cover enough post-pandemic period. The retrieval of hours and other responses by physicians could be motivated by the income losses incurred in the first wave of the pandemic. The long-run responses could be different. The accumulation of demands for regular checkups and postponed physician visits during COVID-19, combined with the existing pre-pandemic crisis, require considerable attention in meeting the service needs of Canadians.

The third paper identifies factors driving declining trends in physician work hours, offering supply-side insights. The weighting technique applied to the LFS survey for profession-specific analysis showcases a broader application for similar research settings. The findings highlight that reduced hours among male physicians and the increased share of females in the workforce play comparable roles in the overall decline. The aging of the physician population, especially the increased percentage of physicians aged 65 and above, also contributes to the reduction.

Escalating absence, a full week or more, rates and increased absence durations are also essential to consider in the decline. Parenthood effects differ across genders, with female physicians working fewer hours if they have children, while male physicians increase their hours upon parenthood. A high variation in the early years of study across regions shows a convergence in the later years due to a higher reduction of hours by regions with initially high average work hours.

The findings of this paper emphasize that changing labour supply patterns by physicians, male-female decomposition, the growing age of the physician population, and the increased absence rate and duration are essential factors to consider in future service provision. Taking them into account, along with the population aging and accumulated demands from COVID-19, can be instrumental in crafting a sustainable strategy for physician workforce planning.