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TAX INCENTIVES AND OLDER WORKERS: EVIDENCE FROM CANADA

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Tax Incentives and Older Workers: Evidence from Canada[*]

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Abstract/Résumé

We provide empirical evidence on the effectiveness of a tax measure aimed at increasing the employment rates of older workers in Quebec, Canada. We use several data sources and various identification strategies. First, we use a Quebec-Ontario difference-in-differences design and do not detect robust effects on employment for most age groups except for those aged 60 to 64, but the common trend assumption is found not to hold. For this last group, we use an alternative identification strategy that exploits the variation in treatment intensity over time using longitudinal administrative tax data for Quebec only. Doing so, we do not find any effect on transitions in or out of the labour force. We do find a small positive effect on earnings (intensive margin) but a negative one on the affected workers' net tax liability. Finally, addressing the invalid comparison with Ontario, we investigate the impact of the credit using a staggered adoption design exploiting differences across cohorts within Quebec. The results are consistent with the alternative approach. We conclude that the tax measure does not appear to be a cost-effective way of raising public revenues nor of increasing the employment rates of older workers.

Nous fournissons des preuves empiriques sur l'efficacité d'une mesure fiscale visant à augmenter les taux d'emploi des travailleurs plus âgés au Québec, Canada. Nous utilisons plusieurs sources de données et différentes stratégies d'identification. Tout d'abord, nous utilisons méthode des différences-en-différences entre le Québec et l'Ontario et ne détectons pas d'effets robustes sur l'emploi pour la plupart des groupes d'âge, à l'exception de ceux âgés de 60 à 64 ans. Cependant, l'hypothèse de tendance commune ne semble pas être vérifiée. Pour ce dernier groupe, nous utilisons une stratégie d'identification alternative qui exploite la variation de l'intensité du traitement au fil du temps en utilisant les données administratives longitudinales sur les impôts pour le Québec seulement. Ce faisant, nous ne trouvons aucun effet sur les transitions dans ou hors de la main-d'œuvre. Nous constatons néanmoins un léger effet positif sur les revenus (marge intensive) mais un effet négatif sur la charge fiscale nette des travailleurs affectés. Enfin,

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pour remédier à la comparaison invalide avec l'Ontario, nous étudions l'impact du crédit en utilisant une conception d'adoption échelonnée, exploitant les différences entre les cohortes au sein du Québec. Les résultats sont cohérents avec l'approche alternative. Nous concluons que la mesure fiscale ne semble pas être une manière rentable d'augmenter les revenus publics ni d'augmenter les taux d'emploi des travailleurs plus âgés.

Keywords/Mots-clés: older workers, labour market participation, tax incentives / travailleurs plus âgés, participation au marché du travail, incitations fiscales

JEL Codes/Codes JEL: J14, J16, H31

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1 Introduction

Population aging in Canada is a major concern. Indeed, while the demographic dependency ratio (DDR) was less 20.4% in 2000, it has reached 31.7% in 2022 and is expected to peak at 59.5% around the year 2082 (OECD, 2019, 2023). Such an increase can have significant economic consequences. An aging population can hinder economic growth due to its effects on saving, investment, and consumption (Baldwin and Teulings, 2014; Eggertsson et al., 2019), lower labour market participation (Burtless, 2013; National Research Council, 2012), or lower productivity (Feyrer, 2008; Maestas et al., 2016). Moreover, an aging population is likely to require a significant increase in government financing, for example to cover rapidly growing health expenditures. All these effects may jeopardize the ability of governments to provide expected services (Lu et al., 2005; Achou et al., 2021; St-Maurice et al., 2021; Clavet et al., 2021). The situation is particularly problematic in several Canadian provinces, notably eastern provinces including Quebec, due to its older population compared to the rest of Canada and a more pronounced preference towards early retirement (Cloutier and Dorion, 2010; Dorion, 2011).

In order to curb the foreseeable consequences of population aging, most governments have sought to increase the employment rates of older workers (OECD, 2006). For Quebec, Bissonnette et al. (2016) show that this could help mitigate the negative effect of population aging on the growth of the labour force and therefore tax revenue. Policies can target monetary incentives or other levers, in particular in the workplace (OECD, 2020). For instance, there is substantial evidence that older workers respond to pension incentives (Börsch-Supan and Coile, 2021) and that their wage elasticity may be large (Blundell and Macurdy, 1999; Goda et al., 2009; Hudomiet et al., 2021). Older workers have been found to be sensitive to provisions of the retirement system such as the removal of earnings tests (Baker and Benjamin, 1999; Friedberg, 2000; Haider and Loughran, 2008).

The retirement income system in Canada provides limited work disincentives for older workers (Milligan and Schirle, 2008). Indeed, the Canada Pension Plan (and its equivalent in Quebec, the Quebec Pension Plan or QPP) is close to actuarially fair for the average worker. In addition, most provisions that penalized or prohibited work while receiving a pension have gradually been removed. In order to foster employment among older workers, governments in Canada need to look elsewhere in their policy toolkit. Although several countries have experimented with payroll tax subsidies with mixed success (Boockmann et al., 2012; Albanese and Cockx, 2019), few have used income tax credits. The only available evidence comes from Australia (Carter and Breunig, 2019) and Sweden (Laun, 2017), and suggests it has had limited impact on employment rates. Consequently, the tax expenditure per additional worker induced to remain in the labour force may be relatively high and may not compensate the additional tax revenue. Hence, the cost-effectiveness of such measures relative to other policy levers is unclear.

In her 2021 mandate letter from Prime Minister Trudeau, Minister of Finance Freeland was given the task *"To address labour shortages and help businesses grow, introduce ... a Career Extension Tax Credit of up to $1,650 a year for seniors who want to stay in the workforce"*. The Federal government has not implemented such a measure as of 2024. Yet, we can learn about the potential effectiveness of such a measure by looking at the experience of provinces who did experiment with such measures. In Canada, the province of Quebec implemented a tax credit in 2012, originally intended for those 65 years of age and older, and later extended to those aged 60 to 64.[1]

The credit increases with earnings and is clawed-back beyond a given ceiling. Between 2012 and 2019, the credit was gradually made more generous. We argue that the gradual expansion of the credit yields sufficient variation to learn something about the effect of tax incentives on the employment of older workers. In Canada, a common identification strategy is to use other provinces as control groups within a difference-in-differences framework. When focusing on specific policies implemented only in Quebec, analysts customarily use Ontario as the comparison group, a province with a similar labour market and industrial composition. Yet, we show, using repeated Labour Force Survey waves, that the identifying assumption underlying this estimator, namely the common trend assumption, is not met in our context.

We then proceed with two alternative identification strategies which exploit differences in eligibility across workers in Quebec as well as the increases in generosity over time. For the sake of robustness, we use two different estimators each with a different dataset. The first estimator uses the Quebec sample of the Longitudinal Administrative Databank (LAD) which contains

[1] The credit more or less corresponds to that mentioned in the mandate letter to Minister Freeland.

a 20% sample of all taxpayers in Canada. We impute the potential credit a worker aged 60-64 could receive based on his earnings and net tax liability when aged 50 to 59. This potential credit, a measure of treatment intensity, is then used to estimate an intent-to-treat effect of the tax credit. The second estimator explicitly accounts for the staggered implementation of the tax credit. It uses the Quebec sample of the Longitudinal and International Study of Adults (LISA) and focuses on a particular dynamic difference-in-differences estimator. The two estimators yield very similar results: the credit is found to have had very little impact on employment. We find small effects on earnings among women, but negative and statistically significant effects on net tax revenue.

In Section 2, we provide institutional details on the tax credit and investigate provincial trends in a standard difference-in-differences setting. In Section 3, we rely on two different identification strategies exploiting within Quebec differences across workers. We discuss policy implications of the results in Section 4.

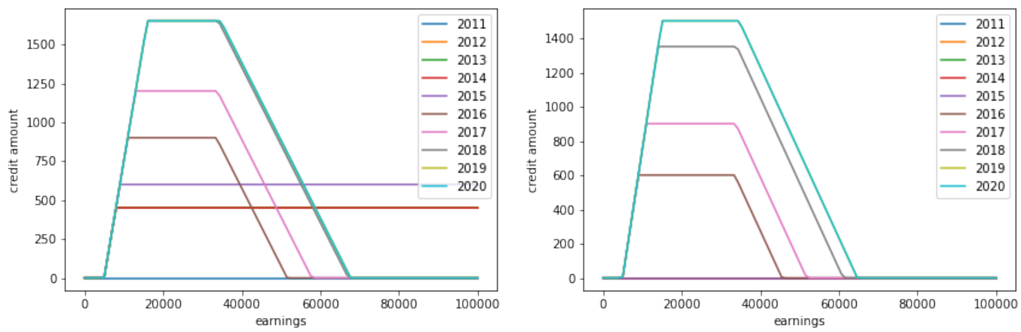
2 Institutional Context and *Prima Facie* Evidence

2.1 The Tax Credit for Career Extension

In 2012, the Quebec government implemented a non-refundable tax credit, the *Crédit pour travailleur expérimenté*, to encourage workers 65 and older to remain employed. The measure was later renamed *Crédit d'impôt pour prolongation de carrière*. We use the acronym `CIPC` henceforth.

When first implemented, workers aged 65 and over were entitled to a tax credit of 15.04% on the first $3,000 exceeding $5,000 of earnings, yielding a maximum credit of $451.20. The ceiling was further increased to $4,000 in 2015 and to $6,000 in 2016, with corresponding maximum credits of $601.60 and $902.40, respectively. However, as of 2016, the credit was clawed back at a rate of 5% on earnings above $33,505. Hence, the credit was entirely clawed back for earnings above $51,553.

Eligibility to the CIPC was extended to those aged 64 in 2016, 63 in 2017, 62 in 2018, and 60 in 2019. While the ceiling was originally set at $4,000 for those aged 64 in 2016, it was increased in 2019 to $10,000 for those 60-64 and to $11,000 for the 65^{+}. The maximum credit in 2019 was thus $1,654.40 for someone earning $16,000 (*i.e.*, $11,000 above the $5,000 threshold). Hence, this represents an 11% wage subsidy for low earners. Figure 1 depicts the changes made throughout the years for the 65^{+} (Figure 1a) and the 64 year-olds (Figure 1b), respectively.



(a) 65 and over

(b) 64 years old

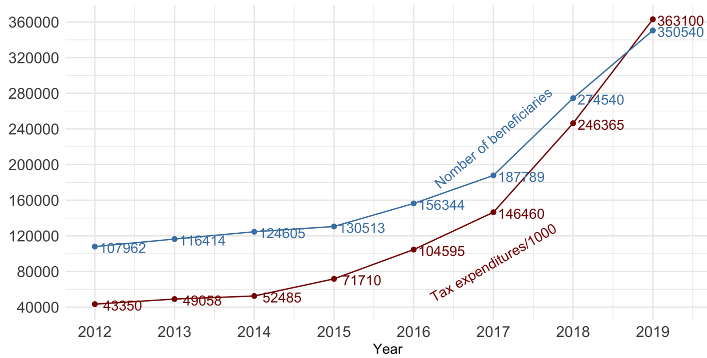
Figure 1: Structure of the Tax Credit for Career Extension (CIPC) by year for the 64 years old and the 65 years old and over.

According to the Quebec Ministry of Finance, the cumulative tax expenditure for the CIPC was $1.2 billion for the years 2014 to 2020, and is projected to reach almost $400 million for 2024 alone. Figure 2 reports the annual number of workers claiming the tax credit as well as the tax expenditures. The Figure illustrates the evolution of the "stock" of workers receiving the CIPC, not the flows into the program. Thus, a worker benefiting from the credit in 2012 and 2013 will contribute to the "stock" in both years. Note there is a significant increase in the "stock" as of year 2016, presumably due to the increased attractiveness of the credit and/or to the lowering of the age of eligibility. All in all, the CIPC is as expensive as the Quebec equivalent of the earned income tax credit for the entire workforce.[2]

2.2 Trends in Employment of Older Workers

As the credit was rolled-out progressively from 2012 onward, it is instructive to analyze the trends in employment rates of older workers over a longer period. It is also useful to contrast the Quebec experience with that of Ontario, a neighbouring province with a similar economic structure. To that end, we use the public-use micro data files (PUMF) from the Labour Force Survey (LFS) for the years 2005 to 2022.

[2]See Tax Expenditures - Individuals (in French).



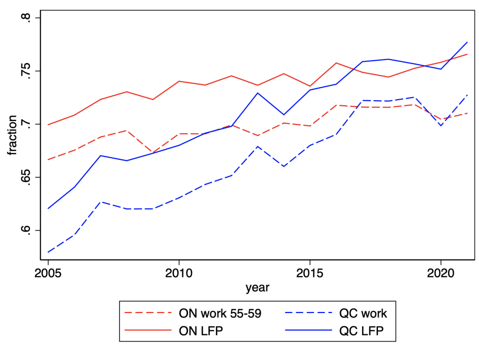
Source: Statistiques fiscales des particuliers
Years 2012—2019

Figure 2: CIPC Yearly Beneficiaries and Tax Expenditures (in thousands of $), 2012–2019.

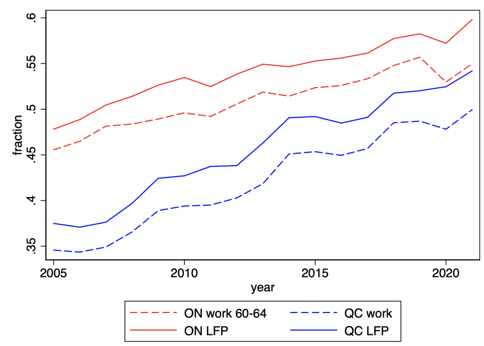
We define labour force participation status as $y_{i,t}$. It takes value 1 if respondent i in year t is active (either working, looking for work or absent from work), and zero otherwise. The PUMF of the LFS include the following age categories, not age *per se*: 55-59, 60 to 64, 65 to 69 and 70+. This is not problematic for those 65 and above since they all face the same credit. It is problematic, however, for those between 60 and 64 as the credit is rolled-out for them from 2016 onward. Those in the 55-59 category serve as an interesting contrast since they are not eligible to the credit. They help assess any cohort trends among younger cohorts occurring while the credit is being implemented. All statistics are weighted using the PUMF-LFS weights.

The employment (activity) rate of age group a in year t and province p is denoted $y_{a,t,p}$. It measures the ratio of the weighted number of employed (active) respondents in a given age group, year and province cell over all respondents in the cell. Figure 3 shows both labour force participation – or activity – and employment rates for different age groups and for Quebec and Ontario separately. All age groups exhibit a positive trend, although there are differences between Quebec and Ontario workers in some cases. In all cases, however, differences between labour force participation and employment rates are constant over time, reflecting a low unemployment rate over the period.

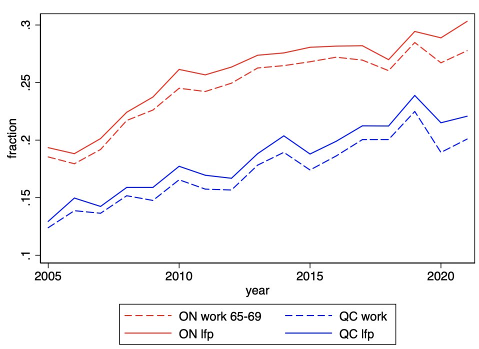
In panel (a), it is readily apparent that the attachment to the labour force of those nearing retirement (55–59) increases in both provinces, with Quebec catching up to Ontario rapidly. Yet, workers in this age group were never eligible to the tax credit. A similar pattern is observed for those in the 60–64 age group (panel (b)). Importantly, the profiles in Quebec



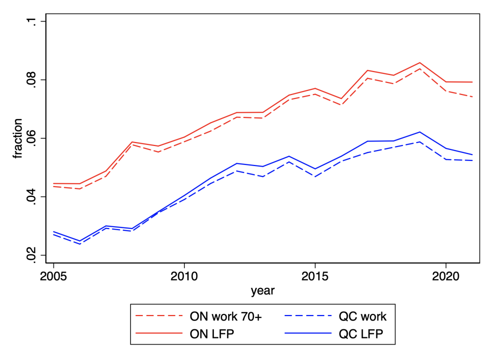
(a) 55-59



(b) 60-64



(c) 65-69



(d) 70+

Figure 3: Labour force participation ("LFP") and employment ("work") rates: LFS data 2005-2022.

are unaltered by the gradual implementation of the CIPC after 2016. Using Ontario as a counterfactual to inform as to what would have happened in Quebec had the credit not been implemented, it must be concluded that the CIPC did not impact the labour force participation or employment rates in Quebec. The same can be said about the 65–69 (panel (c)) and the 70^{+} (panel (d)) age groups, as the gap between Quebec and Ontario remains constant over the CIPC's period of existence, despite the fact that they were all entitled to the credit.

2.3 Difference-in-Differences Estimations

It may be more informative to focus on differences in, rather than levels of, participation and employment rates. These can be investigated using double differences. First, we compute a change in rates, relative to some base year t_0,

$$\Delta y_{a,t,p} = y_{a,t,p} - y_{a,t_0,p}. \tag{1}$$

For those aged 65^{+}, we set $t_0 = 2011$, the year prior to the introduction of the credit, and $t_0 = 2015$ for those aged 60–64, the year prior to their eligibility. The second difference is computed across provinces in order to net out changes that may be due to the CIPC. We compute

$$\Delta^2 y_{a,t} = \Delta y_{a,t,1} - \Delta y_{a,t,0}. \tag{2}$$

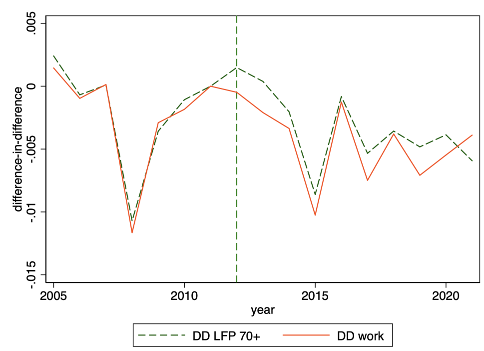
By definition $\Delta^2 y_{a,t_0} = 0$. Numbers are interpreted relative to the difference in rates between Quebec and Ontario in the base year. A positive number implies participation (or employment) rates were increasing faster in Quebec relative to Ontario.

Figure 4 shows the evolution of these double differences for each age group. For those aged 65^{+}, we see very little differential change in Quebec after the implementation of the CIPC. For those 60-64, we see that employment rates in Quebec converge with those of Ontario prior to the introduction of the credit in 2016. Furthermore, we do not see any acceleration after 2016.

(a) 55-59

(b) 60-64

(c) 65-69



(d) 70+

Differences are normalized to zero in 2011 for those 65+ and in 2015 for those 55-59 and 60-64. LFS PUMFS, all rates are weighted.

Figure 4: Difference-in-Differences in Labour Force Participation ("LFP") and Employment ("work") Rates.

We further investigate the differential rates within a regression framework. Let

$$y_{i,t} = \alpha_0 + \sum_{s \neq t_0} \alpha_{1,t} I(t = s) + \alpha_2 qc_i + \sum_{s \neq t_0} \gamma_s I(t = s) qc_i + \epsilon_{a,t}, \quad (3)$$

where $I(q)$ is an indicator function if q is true. It is straightforward to show that without further controls, $\Delta^2 y_{a,t} = \gamma_t$. Define an observation as potentially exposed to the credit using the treatment variable $\text{cipc}_{i,t} = I(t > t_0) \times qc_i$, where t_0 is the year just prior to eligibility. In this case, one obtains the standard difference-in-differences estimating equation,

$$y_{i,t} = \alpha_0 + \sum_{s \neq t_0} \alpha_{1,t} I(t = s) + \alpha_2 qc_i + \gamma \text{cipc}_{i,t} + \epsilon_{a,t}, \quad (4)$$

where we impose $\gamma_t = \gamma$ constant for $t > t_0$ and $\gamma_t = 0$ for $t < t_0$ so that the common trend holds. The parameter γ measures the difference-in-differences estimate of the CIPC. The crucial assumption for this interpretation is that $\gamma_t = 0$ for $t < t_0$, *i.e.* there are no differences in employment trends prior to the implementation of the CIPC. To see this, note that

$$\Delta y_{t,QC} = E(y_{i,t} | p = QC, t) - E(y_{i,t} | p = QC, t_0) = \alpha_{1,t} + \gamma \quad (5)$$

$$\Delta y_{t,ON} = E(y_{i,t} | p = ON, t) - E(y_{i,t} | p = ON, t_0) = \alpha_{1,t}, \quad (6)$$

and therefore that $\Delta^2 y_{a,t} = \Delta y_{t,QC} - \Delta y_{t,ON} = \gamma$.

To estimate equation (4), we use confidential LFS files which provide the exact age so that $\text{cipc}_{i,t}$ can be precisely defined. We control for age even tough the model is estimated for different age groups. We also control for education status as well as the provincial unemployment rate to capture fluctuations in the business cycle. We use labour force participation as the outcome. The model is estimated by least squares and the standard errors are clustered by year and province. Table 1 reports the parameter estimates. For those aged 65^+, we do not find a statistically significant effect of the tax credit. The estimates are 0.0074 for those 65-69 and -0.0092 for those 70+. For the 60-64 group, we find an increase of 2.5 percentage points which is only statistically significant at the 10% level. However, since the common trend assumption is strongly rejected in all the cases, these difference-in-differences estimates are unreliable. In Appendix A, we assess the robustness of these estimates to alternative specifications, involving non-linear estimation methods and various ways of estimating standard errors, as well as sensitivity to using LFS public-use files, not containing single year of age. The conclusions are very similar. Thus, using a standard difference-in-differences design across provinces we find little evidence that the tax credit has increased participation rates among older workers despite Quebec catching up to Ontario.

	Age 70^+	Age 65–69	Age 60–64
QC	-0.0373***	-0.0815***	-0.0791***
	(0.0055)	(0.0064)	(0.0126)
Post(cipc)	0.0356***	0.0304***	0.0402***
	(0.0039)	(0.0067)	(0.0058)
QC× Post(cipc)	-0.0092	0.0074	0.0246
	(0.0069)	(0.0100)	(0.0149)
Age	-0.0155***	-0.0334***	-0.0482***
	(0.0012)	(0.0012)	(0.0015)
Common trend F-stat	126.45	2.39	7.21
(p-value)	<0.001	0.135	0.013

Standard errors in parentheses
*** $p<0.01$, ** $p<0.05$, * $p<0.1$

The table reports the linear probability model estimates of the CIPC DD effect. Standard errors are clustered at the province and year levels (parentheses). Control variables include education and unemployment rates as well as year fixed effects.

Table 1: Difference-in-Differences Estimates of the Effect of the CIPC on Labour Force Participation.

3 Evidence Using Tax Data

There are two potential drawbacks associated with the difference-in-differences approach based on cross-sectional data as used above. First, it omits to account for the fact that eligibility to the CIPC is age- and earnings-dependent.[3] Second, contrary to what was found in Table 1, it assumes that labour force participation in Ontario and Quebec exhibit common trends. To overcome these limitations, we use two different estimators and datasets in what follows in order to investigate the impact of the CIPC on labour market outcomes. The two estimators are based on different sets of identifying restrictions on workers within Quebec.

[3]Because the credit depends on earnings as well as tax liability (it is non-refundable), not all workers are eligible to the credit.

3.1 Intent-to-Treat Using Potential Treatment Intensity

As shown in Figure 1, the CIPC was gradually made more generous. Treatment intensity therefore varies over the period. To exploit this variation, we turn to the Longitudinal Administrative Databank (LAD) which contains a 20% sample of tax filers in Canada (federal tax files). Given this additional variation, and the fact that employment rates in Ontario appear to follow different trends, we focus on within Quebec comparisons. Furthermore, we focus on those aged 60-64 since evidence from Table 1 points to a positive and significant effect at the 10% level (erroneously) assuming that the common trend assumption is satisfied.

The sample used includes individuals aged between 60 and 64 for the years 2013 to 2019. We also restrict the sample to those who have worked at least one year between the ages of 50 and 59 to focus on taxpayers with some labour force attachment. Define $y_{i,t}$ to be employment status of respondent i for year t. Given we observe only taxable earnings and not employment *per se*, we use a threshold of $5,000 to define someone as employed. We focus on transitions in and out of employment between ages 60 and 64. Hence, we construct a sample of those working at $t-1$ and a sample of those not working at $t-1$. The modelling of the variable $y_{i,t}$ on those samples means we are looking at the probability of continuing to work ($y_{i,t} = 1$ in the sample with $y_{i,t-1} = 1$) and the probability of returning to work ($y_{i,t} = 1$ in the sample with $y_{i,t-1} = 0$).

Define $e_{i,t}$ to be earnings of respondent i in year t. Also, define $b_{i,t}$ to be the Quebec tax liability of respondent i in fiscal year t.[4] While we do not have information on take-up of the credit in the data, it can nevertheless by imputed. Indeed, the data in the LAD come from federal tax filings and thus do not contain information on provincial tax measures. But we could compute the credit that someone with earnings $e_{i,t}$ and tax liability $b_{i,t}$ could receive. We could define the potential credit as $\text{cipc}_{i,t} = c_t(e_{i,t}, b_{i,t})$.

Note that the credit function depends on year t as well as earnings and Quebec net tax liability. Given that $y_{i,t} = I(e_{i,t} > 5000)$, running a regression of $y_{i,t}$ on $\text{cipc}_{i,t}$ would yield biased estimates due to obvious simultaneity. We thus construct a measure of potential treatment intensity using earnings and tax liability prior to age 60 (which is the lowest eligibility

[4]The provincial tax liability for the province of Quebec is not observed in the federal tax file. However, Statistics Canada has imputed the value of the tax liability by applying the tax rules to the data available in the federal tax files.

age over the period). This is valid since potential earnings are highly correlated at those ages among those who work. More precisely, we compute $e_{i,0}$ and $b_{i,0}$ as the tax filer mean of these variables between the ages of 50 and 59 conditional on working. [5]

Both variables, $(e_{i,0}, b_{i,0})$, are calculated in constant 2016 dollars, and both are extremely good proxies of potential earnings and tax liability, respectively, to the extent the taxpayer continues to work past age 60, or returns to work. With these values, we calculate a potential *simulated* cipc value given by

$$\text{scipc}_{i,t} = c_t(e_{i,0}, b_{i,0}). \tag{7}$$

Given that $e_{i,0}$ and $b_{i,0}$ are fixed for a given individual, the within variation comes from changes in the credit rules over time. Over the cross-section, this within variation generates variation in treatment intensity that is heterogeneous. We exploit this variation.

Instead of a simulated instrumental IV estimator, we focus on the reduced form,

$$y_{i,t} = \alpha_0 + \alpha_{1,t} + \gamma \text{scipc}_{i,t} + \phi(e_{i,0}, b_{i,0}) + \epsilon_{i,t}, \tag{8}$$

where $\phi(\cdot,\cdot)$ is a flexible spline function of pre-eligibility labour earnings and tax liability (10 knots for earnings and 5 for tax liability).[6] Given the controls, the variation that identifies γ is the variation in the amount of the credit across tax filers over time. As we have shown in Section 2, there is substantial change in the amount but also in the shape of the credit. We scale the amount of the credit, $\text{scipc}_{i,t}$, in thousands of dollars so that a value of 1 indicates a simulated credit of \$1,000. We cluster standard errors at the tax payer level. This is an intent-to-treat estimator. Table 2 reports estimates for men and women separately for both transitions to employment (from not working and from working in the previous period, respectively. We do not find evidence of an effect of the credit on transitions at these ages. The effects are very small, and statistically insignificant. For example, the probability of

[5]Specifically, we compute $e_{i,0} = \frac{\sum_{a=50}^{59}[e_{i,a} \cdot I(e_{i,a}>0)]}{\sum_{a=50}^{59}(I(e_{i,a}>0)}$ and $b_{i,0} = \frac{\sum_{a=50}^{59}[b_{i,t} \cdot I(e_{i,t}>0)]}{\sum_{a=50}^{59}(I(e_{i,t}>0)}$

[6]A simulated instrumental IV approach would have estimated the effect of $\text{cipc}_{i,t}$ by instrumenting $\text{cipc}_{i,t}$ with $\text{scipc}_{i,t}$. Given the mechanical nature of the imputation for $\text{cipc}_{i,t}$ (there would be no variation in take-up given eligibility) we prefer to use the reduced form specification, directly estimating the effect of $\text{scipc}_{i,t}$.

continuing to work increases by 0.3 percentage point when potentially eligible to a credit of $1,000. We compute the counterfactual number of working taxpayers in the absence of the credit and find that it would only decrease the number of taxpayers at work at these ages by 780 (out of 292,000 workers). The estimate for the probability of returning to work is actually negative but again largely economically and statistically insignificant. This holds for both men and women.

	All		Women		Men	
Transition	Return to work	Keep working	Return to work	Keep working	Return to work	Keep working
a. OLS coefficients						
Simulated Credit	-0.0019	0.0039	-0.0018	0.0059	-0.0012	0.0049
	(0.0022)	(0.0029)	(0.0029)	(0.0038)	(0.0036)	(0.0032)
Age (ref: 60)						
61	-0.0125***	-0.0443***	-0.0104***	-0.0535***	-0.0148***	-0.0366***
	(0.0024)	(0.0041)	(0.0027)	(0.0058)	(0.0034)	(0.0039)
62	-0.0195***	-0.0245***	-0.0145***	-0.0255***	-0.0248***	-0.0247***
	(0.0029)	(0.0050)	(0.0031)	(0.0064)	(0.0047)	(0.0052)
63	-0.0263***	-0.0308***	-0.0212***	-0.0375***	-0.0317***	-0.0265***
	(0.0034)	(0.0063)	(0.0035)	(0.0083)	(0.0057)	(0.0064)
64	-0.0323***	-0.0429***	-0.0250***	-0.0494***	-0.0404***	-0.0395***
	(0.0042)	(0.0078)	(0.0044)	(0.0094)	(0.0062)	(0.0083)
Unemployment rate	-0.0054***	-0.0084***	-0.0032	-0.0070*	-0.0081***	-0.0086**
	(0.0016)	(0.0031)	(0.0020)	(0.0041)	(0.0027)	(0.0034)
Observations	252,200	396,400	136,800	181,600	115,400	215,000
R^2	0.0111	0.0093	0.0111	0.0122	0.0151	0.0095
b. Predicted employment with and without credit						
Predicted Employ.	12,795	292,490	5,950	133,110	6,840	159,465
Predicted Employ., credit=0	12,985	291,710	6,055	132,500	6,885	15,8995
Difference	-190	780	-105	610	-45	470

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: We report results separately for six (3x2) groups: for the whole sample, for men and for women; and by employment transition (return to work or keep working) for each. The first panel (a) reports coefficients for the simulated credit as well as for dummies of age and for unemployment rate. The estimation also controls for the year of birth (dummies), a 10-piece spline (knots are at the deciles of mean incomes) of mean income (means of positive income between 50 and 59 years old), and a 5-piece spline (knots at quintiles of mean tax liability) of tax liability (means of tax liability when positive income between 50 and 59 years old). The second panel (b) reports the predicted number of employed workers with and without the credit. Standard errors in parentheses.

Table 2: Estimates of the Effect of the Credit (CIPC) on Employment Transitions.

In Table 3, we run a number of robustness checks. First, we check whether the pre-eligibility period matters. For this, we restrict the computation of average earnings and tax liability from age 50 to 54 instead of 50 to 59. Second, we expand the number of years included in estimation by going back to 2010 instead of 2013. The estimates from the three specifications are very similar to the main ones reported in Table 2.

	All		Women		Men	
Transition	Back to work	Keep working	Back to work	Keep working	Back to work	Keep working
Simulated Credit						
Specification 1 (baseline)	-0.0019	0.0039	-0.0018	0.0059	-0.0012	0.0049
50-59 years old, 2013-2019	(0.0022)	(0.0029)	(0.0029)	(0.0038)	(0.0036)	(0.0032)
Specification 2	-0.0021	0.0023	0.0000	0.0042	-0.0045	0.0026
50-54 years old, 2013-2019	(0.0025)	(0.0026)	(0.0027)	(0.0032)	(0.0042)	(0.0031)
Specification 3	0.0004	0.0031	-0.0004	0.0047	0.0011	0.0047
50-59 years old, 2010-2019	(0.0022)	(0.0028)	(0.0035)	(0.0031)	(0.0027)	(0.0036)

Note: We report the coefficient of the simulated credit for 3 specifications. We report results separately for six separate groups: We report results separately for the whole sample, men and women and by previous status of employment (Employed and Not Employed) . All the estimations control for the age (dummies), the year of birth (dummies) a 10-piece spline (knots are at the deciles of mean incomes) of mean income (means of positive income between 50 and 54 or 59 years old) and of a 5-piece spline (knots at quintiles of mean tax liability) of tax liability (means of tax liability when positive income between 50 and 54 or 59 years old). The first two specification use values from 2013 to 2019 and the third specification extend the period from 2010 to 2019. Standard errors in parentheses.

Table 3: Robustness of the Estimates to the Time Window.

Next, we investigate potential effects at the intensive margin. Since the LAD does not contain data on hours worked, we rely on labour earnings. We estimate the same specification with labour earnings as the dependent variable (in thousands of dollars) and report the results in Table 4. Results in the first three columns reveal a positive effect on earnings driven by women. A one thousand dollar increase in the credit yields an increase of $1,755 in earnings. Is this increase in earnings enough to compensate the cost of the tax measure? We investigate this by running the same regressions as before but use the net tax liability as the endogenous variable. Importantly, the actual tax credit is excluded from the tax liability. Hence the effect of being eligible to the credit on the actual net tax liability is a good measure of the cost-effectiveness of the measure. In the last three columns, we show that the response is negative implying that the tax expenditure does not *pay for itself*, *i.e.* it represents a net tax revenue loss. For women, the net loss is $338 on average. For men, it is $734. Presumably, the negative effect is larger for men since they do not respond at the intensive margin.

	Labour Income			Net Tax Liability		
	All	Women	Men	All	Women	Men
Simulated Credit	1.997**	1.755**	0.649	-0.583***	-0.338***	-0.734***
	(0.814)	(0.858)	(0.863)	(0.109)	(0.085)	(0.130)
Age (ref: 60)						
61	-4.778***	-3.744***	-4.270***	0.098***	0.116***	0.080*
	(0.524)	(0.451)	(0.469)	(0.030)	(0.029)	(0.041)
62	-7.990***	-5.927***	-7.236***	0.064	0.111***	0.024
	(0.941)	(0.773)	(0.866)	(0.050)	(0.037)	(0.071)
63	-10.989***	-7.975***	-10.015***	0.096	0.137***	0.062
	(1.334)	(1.068)	(1.244)	(0.089)	(0.049)	(0.143)
64	-13.831***	-9.812***	-12.766***	0.050	0.184***	-0.076
	(1.697)	(1.352)	(1.595)	(0.091)	(0.064)	(0.121)
Unemployment rate	-0.269	0.439	-0.603	0.159**	0.221***	0.134
	(0.932)	(0.870)	(0.874)	(0.074)	(0.062)	(0.086)
Observations	882,800	429,600	453,200	882,800	429,600	453,200
R^2	0.2466	0.2382	0.2363	0.1840	0.3214	0.1402

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: We report results separately for 3 groups: for the whole sample, for men, and for women. Coefficients for the simulated credit as well as for dummies of age and for unemployment rate are reported. The estimation also controls for the year of birth (dummies), a 10-piece spline (knots are at the deciles of mean incomes) of mean income (means of positive income between 50 and 59 years old), and a 5-piece spline (knots at quintiles of mean tax liability) of tax liability (means of tax liability when positive income between 50 and 59 years old).

Table 4: Reduced Form Estimates of the Effect of the Simulated Credit on Labour Income and Net Tax Liability.

3.2 Dynamic Difference-in-Differences (DD) Estimator

As mentioned above, the standard DD estimator is based on the common trends assumption. Another, often overlooked, assumption states that the treatment effect is constant and homogeneous across cohorts. As is well-known, the DD estimator may be biased if the treatment effect is heterogeneous across cohorts and time (Roth et al., 2022; Goodman-Bacon, 2021). This is because traditional DD analyses are partly based on so-called "forbidden" comparisons, *i.e.* comparisons which include individuals who have already been treated in the control group of individuals who become treated. In our context, there are potentially three distinct groups: the never-treated (Ontario), who always act as a control group; the not-yet-treated, who temporarily act as a control group but are destined to be treated; and the treated. This is a not inconsequential variation on the standard design explained above. Indeed, in the standard case, the "not yet treated" group does not exist.

To illustrate how the implementation of the CIPC may lead to a biased standard DD estimator, consider individuals who were 61 years of age in 2008. When the CIPC came into effect in 2012, these individuals were aged 65 and therefore eligible for the credit. Individuals aged 60 in 2008 were only 64 in 2012 and therefore only eligible for the CIPC in 2013. Consequently, the individuals in this cohort belonged to the not-yet-treated group in 2012 only. In 2013, they joined the treated group. The same applies to the younger cohorts, who belong to the not-yet-treated group before eventually joining the treated group.

To fully account for the staggered implementation of the CIPC, we use the estimator proposed by Callaway and Sant'Anna (2021).[7] The main benefit of this estimator is to allow the treatment effect to be disaggregated into subgroups, g – each representing a different cohort made up of individuals who became eligible to the CIPC during the same year and designated by that year – and to allow the treatment effect to vary across time, t. The parameters of interest are written as follows:

$$ATT(g,t) = E[y_{i,t} - y_{i,g-1}|G_i = g] - E[y_{i,t} - y_{i,g-1}|G_i = g'] \text{ for any } g' > t, \tag{9}$$

[7] We have used different estimators to estimate the treatment effect of the CIPC. In particular, we have used the two-step procedure proposed by Butts and Gardner (2022) and Gardner (2022) which allows to easily control for exogenous variables. The different estimators all yield similar results.

where $ATT(\cdot)$ represents the "average treatment effect on the treated". The expression $ATT(g, t)$ designates the effect of the CIPC at period t for those who qualified in cohort/year g. The different cohorts are grouped into subsets $G_g = 1$. There are as many estimators as there are "treated" groups and periods. As with the standard DD estimator, the identification of the estimator in equation (9) requires that the common trend assumption be satisfied. The simplest extension of the common trend assumption to the staggered case requires that it holds for all combinations of periods and all combinations of groups treated at different times.[8] In particular, one variant of the Callaway and Sant'Anna (2021) estimator only requires that the common trend assumption be satisfied for the treatment periods. This may be written as follows:

$$E[y_{i,t}(\infty) - y_{i,t'}(\infty)|G_i = g] = E[y_{i,t}(\infty) - y_{i,t'}(\infty)|G_i = g'], \text{ for all } t, t' \geq g_{min}, \tag{10}$$

where g_{min} corresponds to the first period a unit is treated (*i.e.* 2012 or later), and $y_{i,t}(\infty)$ is the potential outcome in the absence of the treatment.

To implement the Callaway and Sant'Anna (2021) estimator, our analysis uses data from Statistics Canada's Longitudinal and International Study of Adults (LISA). LISA is a biennial longitudinal study that has been produced since 2012 and collects information on education, training, family, health, income and employment. In addition, each wave includes historical administrative data from the T1 Personal Master File, T1 Family File, T4 Summary and Supplementary Files, and Pension Plan. We use all five available waves: 2012, 2014, 2016, 2018 and 2020. The data for each wave relates to the previous tax year. Individuals are given a unique identifier so that a panel can be assembled. Our sample covers tax years 1981–2019. The panel thus contains tax information before and after each wave. We use those data instead of the LAD because we want to control for other covariates, most notably education, that are not available in the LAD.

For our purpose, we must restrict the LISA sample in order to match it closely to our population of interest. To complement the analysis of the previous section, we limit our sample to individuals aged between 55 and 72 during the 2008–2019 period. After careful analysis, it

[8]Note that the estimator assumes that the treatment has no causal effect before its implementation (no anticipation). Otherwise, the estimator of the treatment effect will inevitably be biased upwards or downwards. The estimator proposed by Callaway and Sant'Anna (2021) admits a limited form of anticipation.

was decided to limit the pre-CPIC period to the years 2008–2011. This gives us a pre-CPIC window of four years and a post-CPIC introduction window of seven years. Furthermore, our sample is limited to the residents of Quebec for the reason mentioned above. The time frame allows us to consider as many as 6 treatment groups (2012-2017) and 8 time periods (2012-2019). Reporting that many $\widehat{ATT}(g,t)$ parameters is cumbersome, and some may be imprecisely estimated. Instead, we report the estimates for k periods after the treatment was initiated across the different cohorts. This is computed as:

$$ATT_k^{\omega} = \sum_g \omega_g ATT(g, g+k), \tag{11}$$

where the weights, ω_g, can be chosen to weight different cohorts equally or according to their relative frequencies (see Callaway and Sant'Anna, 2021 for details).

Table 5 provides a series of $ATT(t)$ estimates for $g = 2012, \ldots, 2017$ and $t \geq g$.[9] The parameter estimates are provided for the entire sample, as well as for women and men separately. We also distinguish between the extensive margin (work/no work) and intensive margin (earnings).[10] The first noteworthy feature of the table is that few parameter estimates are statistically significant. When considering the entire sample, the only parameter estimates at the extensive margin that are statistically significant bear a negative sign.[11] At the intensive margin, on the other hand, we do observe a positive impact for the years 2018 and 2019. This result is entirely driven by women. Indeed, while they do not respond to the CIPC at the extensive margin, their labour earnings increase marginally as of year 2016, *i.e.* when the CIPC was made more generous and accessible to younger workers. Men, on the other hand, do not respond either at the extensive nor at the intensive margins. If anything, their responses to the CIPC at the extensive margin is surprisingly negative during the years 2014-2016.

All in all, the empirical evidence brought to bear using a staggered estimator lends little support to the effectiveness of the tax credit on employment. Only when the CIPC is made most generous do we observe a small impact at the intensive margin among women. The results are consistent with those found using the empirical strategy of the previous section.

[9] All the regressions control for age, education and unemployment rate as in the previous section.

[10] Employment is defined as labour income above \$5,000 as in the previous section. The results are robust to variations in this threshold.

[11] Standard errors are clustered at the respondent level.

Year	Entire Sample		Women		Men	
	Work	Earnings	Work	Earnings	Work	Earnings
2012	-0.001	-1080.630**	0.019	738.808***	-0.022	-2907.282***
	(0.015)	(541.676)	(0.019)	(274.565)	(0.024)	(1039.317)
2013	-0.028*	-698.337	-0.017	706.669	-0.042	-2123.268**
	(0.017)	(609.652)	(0.022)	(612.474)	(0.027)	(1056.406)
2014	-0.034*	-1038.028	-0.006	337.295	-0.067**	-2488.736**
	(0.020)	(683.846)	(0.026)	(647.018)	(0.030)	(1222.404)
2015	-0.045**	-284.442	-0.009	1124.494	-0.094***	-1785.434
	(0.020)	(746.592)	(0.029)	(744.954)	(0.031)	(1301.676)
2016	-0.018	976.117	0.025	1407.600**	-0.044*	506.009
	(0.018)	(640.148)	(0.033)	(647.532)	(0.027)	(1102.812)
2017	-0.020	840.510	0.001	1135.721*	-0.040	499.853
	(0.018)	(672.306)	(0.038)	(624.801)	(0.027)	(1207.481)
2018	0.012	2710.267**	0.048	1836.961**	0.001	3550.115
	(0.020)	(1407.712)	(0.041)	(790.294)	(0.029)	(2810.279)
2019	0.019	2457.110**	0.066	2379.642***	0.034	2413.722
	(0.023)	(1019.951)	(0.047)	(932.455)	(0.034)	(1846.156)

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 5: Year-Specific CIPC Treatment Effects.

4 Conclusion

In this paper, we study the effects on employment of older workers of offering a tax credit for earned income. In 2012, the province of Quebec started offering such a credit. The credit was made more generous in 2016. Cumulative spending on this tax measure since inception is over one billion dollars. One of the aims of the measure was to increase tax revenue by increasing attachment to the labour force. Studying the effectiveness of this measure is important given that the Federal government announced its intention of offering a similar measure to Canadians in the 2021 Mandate letter for the Minister of Finance.

We use a variety of approaches to study the effect of the measure on employment. First, we use a difference-in-differences strategy by comparing the evolution of employment in Quebec to the evolution in Ontario, a province that did not implement such a measure. We find no evidence of an increase in employment for those 65 and over, and limited evidence prior to age 65. We reject the common trend assumption for such comparisons. Second, we use the Longitudinal Administrative Databank to study transitions in and out of the labour force between the ages of 60 and 65. We exploit variation in the generosity of the credit over time to identify its effect. Contrary to what was found using a standard DD approach, we find no evidence that the CIPC has had any effect on employment. We do find some evidence of a positive effect on earnings among women only. However, their overall net tax liability is negative. Hence, the CIPC does not contribute to raising tax revenues. Finally, we investigate the robustness of difference-in-differences comparisons using a dynamic treatment effects approach. The results are consistent with other approach.

What could explain this lack of effectiveness? There are a number of reasons. First, it may be that the tax measure is not well-known. Indeed, the tax credit is non-refundable and only salient when filling a tax return. A second possibility is that it is simply not generous enough to induce workers to extend their career given their preference for leisure, which probably increases with age. One must consider that pension considerations may be more salient to the retirement decision than tax considerations. Third, it may be that some workers could be inclined to extend their career part-time given the strong clawback of the measure, but may be prevented from doing so due to inflexible work arrangements in their current occupation. This type of rigidity may hinder the effectiveness of the tax measure.

Our analysis does not go beyond 2019 due to our identification strategy. The pandemic makes any analysis beyond 2019 more hazardous as it impacted provinces differently for various reasons. Nevertheless, it is interesting to look at the trends in employment rates in recent years among older workers. Figure 5 reports monthly employment rates for those between 60–64 years of age from 2017 to mid-2024. One can see that Quebec has converged with other large provinces by 2024. Since 2023, employment rates have reached a plateau in Quebec similar to that in other provinces. We have shown that Quebec's convergence likely has little to do with the CIPC tax credit. One could reasonably assume that it is very unlikely a new tax measure at the federal level, with similar features and generosity, could boost employment rates even higher without substantial costs to other taxpayers.

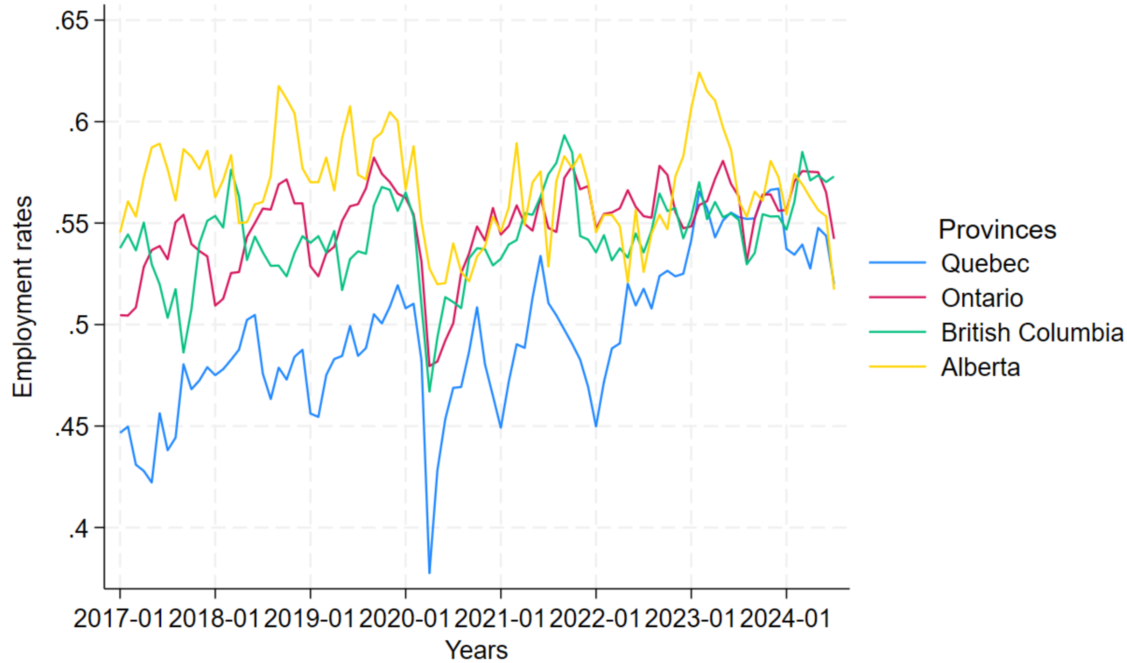


Figure 5: Monthly employment rates, age 60-64: LFS data 2017-2024.

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A Provincial Difference-in-Differences Estimates

In Table 1, we report difference-in-differences estimates using the detailed (confidential) LFS files and linear probability models. In this appendix, we show how these estimates differ if we use public files (PUMF) instead, which do not contain single years of age; different standard errors; and non-linear – probit – models for the difference-in-differences. For example, Cousineau and Tircher (2021) estimate a probit model on PUMF LFS data using the following difference-in-differences specification:

$$y_{i,t} = I(\alpha_0 + \alpha_1 post_t + \alpha_2 qc_i + +\gamma \text{cipc}_{i,t} + \mathbf{x}_{i,t}\beta + \epsilon_{i,t} > 0) \tag{12}$$

where $post_t = I(t > t_0)$ and $\mathbf{x}_{i,t}$ include a set of controls, namely education and the province-specific unemployment rate. The error term $\epsilon_{i,t}$ is assumed iid and normal leading to a probit formulation. The treatment variable is now $\text{cipc}_{i,t} = qc_i \times post_t$. These restrictions in year effects imply no pre-trends in employment rates and a constant jump in the *post* period ($\alpha_{1,t} = 0$ for $t <= t_0$ and $\alpha_{1,t} = \alpha_1$ for $t > t_0$). Cousineau and Tircher (2021) interpret γ as the difference-in-differences estimate of the effect of the credit. They estimate this specification separately by age group (60-64, 65-69 and 70+). In column 1 of Table A1, we report marginal effect estimates (at the mean) which are extremely close to those reported in their study. In the case of the effect of $\text{cipc}_{i,t}$ the authors compute the marginal effect at the mean of the other variables. Denote $\overline{z} = (1, \overline{post}_t, \overline{qc}_i, \overline{\mathbf{x}}_{i,t})$ the vector of means and $\overline{z}'\eta$ where $\eta = (\alpha_0, \alpha_1, \alpha_2, \beta)$. Given the specification above this amounts to calculating

$$E(y_{i,t}|\overline{z}, \text{cipc}_{i,t} = 1) - E(y_{i,t}|\overline{z}, \text{cipc}_{i,t} = 0) = \Phi(\overline{z}'\eta + \gamma) - \Phi(\overline{z}'\eta) \tag{13}$$

where $\Phi(\cdot)$ is the normal CDF. This marginal effect does not have a difference-in-differences interpretation. The calculation needs to account for the fact that $\text{cipc}_{i,t}$ is an interaction of two variables already in the model (*qc* and *post*) (Ai and Norton, 2003). Omitting the dependence on x, the correct marginal effect of the interaction term, which amounts to a difference-in-differences estimate, is

$$\Delta^2 y = (\Phi(\alpha_0 + \alpha_1 + \alpha_2 + \gamma) - \Phi(\alpha_0 + \alpha_2)) - (\Phi(\alpha_0 + \alpha_1) - \Phi(\alpha_0)) \tag{14}$$

Note that the Quebec time difference (the first term in brackets) is evaluated at $\alpha_0 + \alpha_2$ while the difference for Ontario is evaluated at α_0. Given the fact that $\Phi(\cdot)$ is non-linear, one could obtain a positive γ and a negative $\Delta^2 y$. We report these alternative marginal effect estimates of the CIPC in panel *b.* of Table A1, using also an overall sample mean for the x, which does not influence the estimates much. We now obtain that the sign of the marginal effect for the 70+ group changes. The CIPC would have had a negative effect on participation for that group. For the 65-69 group, the reported effect is cut by half. For the group 60 to 64, the non-linear transformation has only a minor effect on the estimate.

The non-linear difference-in-differences depends on the values of x used. These change over time. Most studies in this literature move to a linear probability model to compute these differences in differences. Marginal effects are often very similar although there is often a loss of precision. Row 1 of panel *c.* in Table A1 reports that these marginal effect estimates are the same as one would obtain if we estimated the specification using a linear probability model (OLS). Hence, going forward, we focus on linear probability model estimates. These estimates are also very similar to those reported in Table 1 which suggests that measurement error introduced by not having single years of age has little effect on the estimates.

One important problem when estimating a difference-in-differences equation over two groups is that one grossly overestimates precision by assuming observations are iid. In fact, one could aggregate the data by year and province and run the same specification. One can account for the fact that some cells have more observations by weighting those cells by the number of observations. When we do this, standard errors are much larger and no estimate is statistically significant. We can also cluster standard errors to account for correlation across observations. However, we only have two provinces which precludes from allowing for correlation over time for the same province. Instead, we can cluster at the province and year levels which, unsurprisingly, yields very similar estimates to those obtained by collapsing the data. The last and most problematic issue with these estimates is that the common trend assumption ($\gamma_{1,t} = 0$ for $t < t_0$) does not hold for two of the three age groups (panel *d.* of Table A1). This result is very similar to that reported in Table 1.

	age 70+		age 65-69		age 60-64	
	(1)	(2)	(3)	(4)	(5)	(6)
a. Probit marginal effects						
QC	-0.026***	-0.024***	-0.089***	-0.089***	-0.085***	-0.102***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
post	0.018***	0.016***	0.028***	0.022***	0.034***	0.052***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
QC × post (cipc)	0.003**	-0.001	0.016***	0.006**	0.028***	0.036***
	(0.001)	(0.001)	(0.003)	(0.003)	(0.003)	(0.003)
unemployment rate		0.012***		0.027***		0.043***
		(0.001)		(0.001)		(0.001)
university		0.074***		0.129***		0.133***
		(0.001)		(0.002)		(0.002)
some college		0.042***		0.084***		0.101***
		(0.001)		(0.002)		(0.002)
high school		0.030***		0.059***		0.075***
		(0.001)		(0.002)		(0.002)
rrq		0.018***		0.046***		0.057***
		(0.001)		(0.002)		(0.002)
Observations	928,969	928,969	460,365	460,365	558,830	558,830
b. Non-linear marginal effect						
marginal eff. DD	-0.003**		0.007**		0.0271**	
	(0.001)		(0.003)		(0.003)	
c. LPM estimates with SE.						
no clustering	-0.003***		0.007***		0.027***	
	(0.001)		(0.003)		(0.003)	
cluster yr x prov	-0.003		0.007		0.027	
	(0.004)		(0.009)		(0.0163)	
collapse yr x prov	-0.003		0.007		0.0271	
	(0.005)		(0.0127)		(0.020)	
d. Common trend T-stat		36.24		0.65		14.64
(p-value)		<0.001		0.428		<0.001
e. Confidential LFS						
Individual year of ages	−0.0092		0.0074		0.0246	
	(0.0069)		(0.0100)		(0.0149)	

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table A1: Difference-in-Differences Estimates of the Effect of the Credit on Labour Force Participation. We report results separately for three groups: ages 70+ (columns 1 and 2), ages 65-69 (columns 3-4) and ages 60-64 (columns 5-6). The first panel (*a.*) reports marginal effects at the mean for two specifications by age group, one without controls and one with the same controls as Cousineau and Tircher (2021). Panel *b.* reports the non-linear marginal effects accounting for the interaction effect in the DD specification. Panel *c.* reports linear probability model estimates of the DD effect with different types of standard errors. It first reports estimates without clustering, then assuming clustering at the province and year levels; and finally we report results for a regression on aggregated data by year and province, where each observation is weighted by the number of observations in each cell. Panel *d.* reports the t-test statistic (along with p-value) on a pre-treatment linear trend interaction with a Quebec indicator. Finally, the last panel (*e.*) reports the estimation of the linear probability model with clustering at the province and year levels estimated on confidential LFS data that include the individual year of age. Standard errors in parentheses.