The CoViD-19 Pandemic and Mental Health: Exploring Crucial Channels

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Abstract

This paper studies the co-evolution of mental health with some of the most prominent risk factors associated with the CoViD-19 pandemic for the working population. In order to do so, we exploit data from the Dutch LISS panel, which contains information from before the pandemic and five customized questionnaires during its first year. Mental health decreased sharply with the onset of the first lockdown but recovered fairly quickly. At the end of 2020, levels of mental health are comparable to those a year earlier. At the individual level, we show that perceived risk of infection, labor market uncertainty, and emotional loneliness are all associated with worsening mental health. Both the initial drop and subsequent recovery are larger for parents of children below the age of 12. Among parents, the patterns are particularly pronounced for fathers if they shoulder the bulk of additional care. Mothers' mental health takes a particularly steep hit if they work from home and their partner is designated to take care during the additional hours.

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1 Introduction

Starting in early 2020, the CoViD-19 pandemic and its ramifications upended the lives of billions of people. From early on, researchers and practitioners pointed towards possible adverse effects on population mental health through a variety of channels. Some of the most prominent pathways identified by prior literature include worries about and occurrence of the health effects of contracting the virus (e.g., Hollingue et al., 2020; Kämpfen et al., 2020); increased loneliness through the loss of social contacts (Etheridge and Spantig, 2022); anxiety about job and income losses in the wake of the global recession caused by the pandemic (e.g. Davillas and Jones, 2021; Kämpfen et al., 2020; Witteveen and Velthorst, 2020); and increased stress in families with children affected by closures of schools and daycare (Etheridge and Spantig, 2022; Zamarro and Prados, 2021).

This paper studies the joint evolution of mental health and these four risk factors at the individual level. That is, we exploit information measured at several points in time—both before the pandemic and during its first year—in a fixed effects framework. We focus on the working population because two of the four channels are less relevant for other groups. Our data stem from the LISS panel, which is an Online Panel based on a probability-based sample of the Dutch population. Before the pandemic, panel members participated in regular surveys about their health, time use, and socio-economic factors. During the pandemic, we fielded five customized questionnaires that included information about mental health and the four risk factors. We do all our analyses separately for men and women because the literature has revealed various gendered patterns.

Our first main finding is that on average, mental health dropped sharply right at the onset of the first lockdown in mid-March 2020, which included many types of contact restrictions along with closures of schools and daycare centers. The drops are large and of similar size for both genders. Average scores recovered within a few months; towards the end of the year they were at or above the levels found in prior years during that season. These findings mirror those in related literature (see e.g. Robinson et al., 2022, for an overview). Using nationally representative data from the U.K., Daly et al. (2020), for example, find an increase in the prevalence of mental health problems in April – June 2020 compared to pre-pandemic levels, with clear signs of recovery toward mid-2020. These patterns are consistent with other evidence for the U.K. (e.g. Banks and Xu, 2020; Etheridge and Spantig, 2022; Proto and Quintana-Domeque, 2021) and with similar work for the

U.S. (e.g., Daly and Robinson, 2021). At any point in time, mean levels of psychological distress are higher in women, a result that is well documented in the literature (e.g., Kessler et al., 1993; Van de Velde et al., 2010). This is also true in our data. However, in contrast to, for example, the U.K. (Etheridge and Spantig, 2022), we cannot exclude the possibility that the pattern of average mental health scores in the Netherlands evolves as a parallel shift between working men and women over our observation period.

Second, we show that all four channels are associated with changes in mental health. This holds true even when considering all risk factors jointly at the individual level, conditioning on pre-pandemic levels. To the best of our knowledge, this is among the first papers to have done so. Similar to, for example, Hollingue et al. (2020) and Kämpfen et al. (2020), we find that subjective infection risk is negatively associated with changes in mental health; effects are significant only in the female sample. Increases in loneliness are associated with lower mental health; the size of the coefficient is substantially larger for women which is in line with findings for the U.K. (see e.g. Etheridge and Spantig, 2022). As in other recessions, perceived labour market risk is associated with worse mental health also during the pandemic (see e.g. Davillas and Jones, 2021; Kämpfen et al., 2020; Witteveen and Velthorst, 2020). We find both subjective unemployment risk and large reductions in the hours worked—a useful measure for labour market risk in the presence of firing restrictions and employment protection programmes (Zimpelmann et al., 2021)—to be strongly associated with mental health problems for men. Finally, parents of young children experienced larger drops in mental health during the initial lockdown, a result which has been similarly shown in other studies (see e.g. Etheridge and Spantig, 2022; Zamarro and Prados, 2021).

Third, we demonstrate important heterogeneity in parents' mental health changes. If parents shared the extra care duties created by the closures of schools and daycare facilities, the initial drop in the mental health score was small if present at all. In contrast, if only one parent shouldered the additional childcare, that parent has consistently lower scores—relative to before the pandemic—over the entire year. The drop in March was particularly pronounced for fathers who took on the additional duties by themselves. Digging deeper into this, we show that working from home might be an important mechanism for this finding. The negative relation between caregiver duties and mental health is most pronounced for fathers who work many hours from home. Looking at this from a different angle, we exploit time use data from November 2019 and April 2020, which contain a direct measure of the number of hours worked from home while being responsible for children *at the same time*. There are strong gender differences: Mental health is hump-shaped in such hours for men and U-shaped for women. The respective peak/trough is around 15–20 hours. These different patterns are consistent with the fraction of total working hours spent simultaneously on childcare and work. Since men work longer hours, men still have plenty of time to get some work done if they spend 15–20 hours taking care of children while working. For women, this is much less the case. Altogether, our results are consistent with a "mom is never off duty when home"-effect.

In the next section, we describe the Dutch setting, our sample, and the evolution of the key variables over the period from late 2019 and throughout 2020. In Section 3, we first explain our various fixed effects regressions before describing the results. Section 4 discusses our findings and concludes.

2 Context, data, and stylized facts

In this section, we outline the setting for our analysis. We first describe the institutional context in the Netherlands, putting particular emphasis on the temporal evolution of social distancing policies enacted to reduce the spread of SARS-CoV-2. Following that, we provide an overview of the dataset we collected. We then describe the evolution of mental health and the key explanatory variables throughout our period of study between November 2019 and December 2020.

2.1 The CoViD-19 pandemic and social distancing policies in the Netherlands

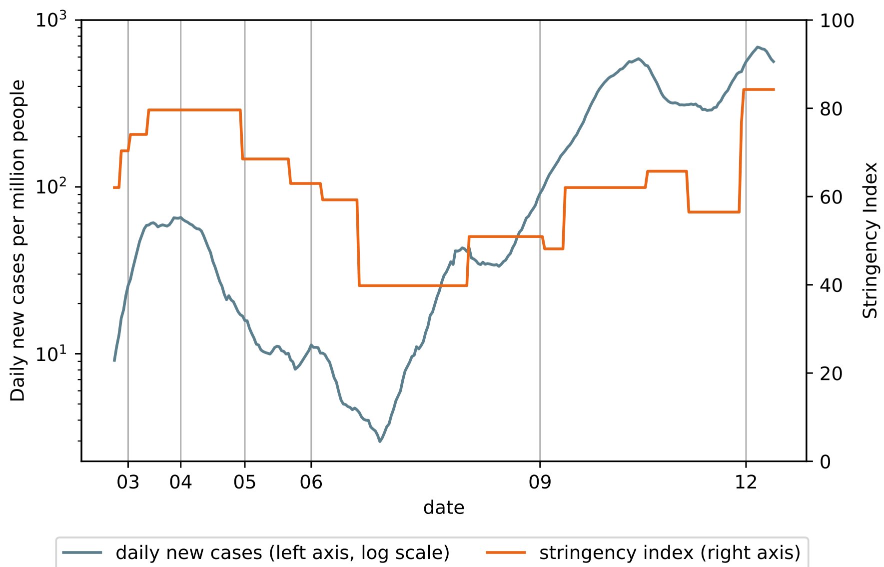
The first SARS-CoV-2 infection was detected in the Netherlands in late February 2020. By mid-March, more than 10 new cases per million inhabitants were confirmed each day (all infection number are based on Roser et al., 2020). Figure 1 shows the number of daily new cases in the Netherlands per million inhabitants on a logarithmic scale (left axis). This number had reached 60 by the end of March—despite limited testing—and stayed roughly at that level for the first three weeks of April. It declined thereafter and between May and late July 2020, new cases were below 10 per million inhabitants and day. Infection rates started to rise again in July. By mid-September, they had surpassed the peak of the first wave, and in late October, they hit almost 600 new daily cases per million inhabitants. During November, this number decreased to a value below 300 but steeply rose again and peaked just below 700 cases before Christmas.

Similar to other countries, the initial rise in infections prompted the Dutch government to impose restrictions on economic and social life to slow down the virus' spread. The Oxford Response Stringency Index (Hale et al., 2020) quantifies the extent of these restrictions. It runs from zero (absence of any restrictions) to one hundred (all included restrictions in full effect). Its evolution from the beginning of our CoViD-19 surveys to the end of the year is shown in Figure 1. In mid-March, all schools and childcare facilities were closed along with restaurants, cafes, bars, and other businesses involving personal contacts. People were advised to stay at home, to keep a distance of at least 1.5 meters to each other, and to avoid social contacts; the number of visitors at home was restricted to a maximum of three individuals. While most of these policy measures resembled those in other European countries, they did not involve a general curfew and some measures were much more lenient. Businesses, such as stores for clothes, utilities, or coffee shops remained open as long as they could guarantee to maintain the social distancing rules. The government advised everybody to stay at home, but people were allowed to go outside without any official permission, and they were allowed to meet a maximum of three other non-household members as long as social distancing was maintained. Public locations were accessible and traveling or the use of public transportation was possible throughout.

Beginning in May, the restrictions were gradually lifted. Daycare facilities and primary schools were among the first areas to open up again, secondary schools followed in early June. With the exception of bans of larger (inside) gatherings, social and economic life was largely back to what it was before the pandemic.

In mid-October, the Dutch government tightened restrictions again in response to the steep rise in infection numbers. Many were stricter than those imposed in March 2020: Besides the closure of restaurants, bars, museums, and other public places, opening hours for shops were limited and the sale of alcohol was prohibited after 8 p.m.. An important difference compared to the restrictions imposed in March–May 2020 was that schools and daycare centers remained open. Along with a temporary sharpening of the measures in early November, this brought infection rates down for some time. However, their rise during the first half of December prompted yet another tightening effective from December 15, at just about the time our data collection ended. All shops except supermarkets and essential services were closed along with childcare facilities and schools which

Figure 1: CoViD-19 cases and Oxford Response Stringency Index in the Netherlands in 2020.



Note: Daily new cases per million inhabitants are shown in blue on the left scale. Source: Roser et al. (2020). The Oxford Response Stringency Index measures the stringency of non-pharmaceutical interventions, see (Hale et al., 2020). Vertical lines indicate the end dates of surveys during the year 2020 (fieldwork was the previous 10 days in March and the previous four weeks in other months)

started the winter break a week earlier.

2.2 Data and sample construction

Our empirical analysis uses the Longitudinal Internet Studies for the Social Sciences (LISS), which is a panel data set based on a probability sample of the Dutch population. The LISS panel has been running since 2007 and comprises roughly 7,000 individuals from about 4,000 households. Each month, respondents are invited to complete questionnaires lasting 15-30 minutes on average. The information solicited from respondents includes a set of ten core questionnaires repeated every year and questionnaires designed by external researchers.

Our baseline measure of mental health stems from the core questionnaire on health administered in November 2019. We included the same measure in a set of modules that we designed to track the consequences of the pandemic (Gaudecker et al., 2020). In these questionnaires, we asked about mental health, labor market outcomes, and expectations during the CoViD-19 crisis. The initial module was fielded in late March 2020, a week into the first lockdown. Five more modules followed in April, May, June, September, and December. All CoViD-19 survey modules were addressed to all panel members aged 16 years and older; response rates exceeded 80% in all waves.

The basic structure of our data is an individual-level panel with up to six time series observations.[1] We make the following restrictions on our sample. We keep household heads and their partners for whom we have at least two complete sets of observations. We restrict the sample to individuals up to age 70 who reported to be employed or self-employed just before the pandemic started while working positive hours. Our resulting sample consists of 10,203 observations on 2,320 individuals. Basic demographic characteristics by gender are shown in Table 1. In line with our sample restrictions, average age is in the upper fourties for both men and women. Close to half of the sample has some form of tertiary degree, reflecting high education levels among younger cohorts.

Table 1: Sample characteristics by gender

	Men	Women
age	47.8	45.9
low education	0.14	0.13
medium education	0.38	0.40
high education	0.48	0.47
individuals	1,119	1,201
observations	5,010	5,193

2.3 Mental health

The core LISS questionnaire on health contains the MHI-5 (Mental Health Inventory 5) measure, which is a brief, validated international instrument for assessing mental health in adults (see, e.g., Berwick et al., 1991; Thorsen et al., 2013). The MHI-5 score is a five-item subscale of the Short Format 36 (SF-36), a comprehensive tool to measure the prevalence of depression. Hoeymans et al. (2004) compare the MHI-5 measure to the General Health Questionnaire (GHQ-12) for the Dutch population. They find both measures to be similarly predictive for mental health problems.

[1] Because of the short time span between the initial wave in late March and the second wave in April, we did not ask about mental health in April.

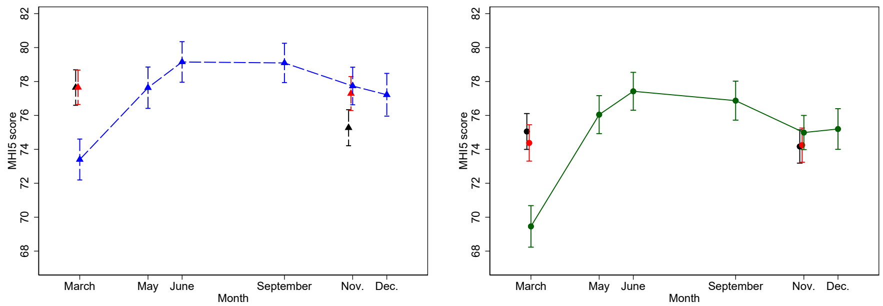
The MHI-5 instrument consists of five separate questions to assess how people felt in the past four weeks (see Online Appendix A for details). Each answer comes on a six-point scale. To obtain the MHI-5 score, all answers are coded on scales from zero to five such that higher values indicate better health. Individual values are summed up and multiplied by four. The resulting MHI-5 score ranges from zero to 100, with zero representing very poor mental health and 100 representing its best possible level. Medical literature generally uses cutoffs between 52 and 76 to dichotomize the measure (e.g., Cuijpers et al., 2009; Hoeymans et al., 2004; Thorsen et al., 2013). Values below the cutoff are interpreted as indicative of mental health problems. Probably the most common cutoff is 60, which is also used by official statistics (Statistics Netherlands, 2015). In our analyses, we use the raw score in order to work with a near-continuous measure; averages are typically in the 65-80 range. Our measure is meaningful in the sense that it is indicative of variation in this critical range (as opposed to measuring changes between nearly optimal and optimal; or variation very close to zero).

Using the continuous measure or cutoff-based indicators derived from it makes the assumption that the MHI-5 score is cardinal and interpersonally comparable. In particular, we require that the scale is the same for men and women. A recent study by Bond and Lang (2019) points out that such a cardinality assumption is not innocuous because it is not robust to monotonic transformations. While that study refers to happiness scales—where the subjective element inherent to the scale is larger than when it comes to quantifying observations about oneself, which are objectively verifiable in principle—it is important to keep in mind when interpreting our results. This said, the assumption is inherent in clinical practice and official statistics, which use a common cutoff regardless of gender.

In the LISS panel, the MHI-5 battery is asked every year in November as part of the core health module. In the past few years, it has also been asked to all panel members as part of the "VICTIMS" study (e.g., Van der Velden et al., 2021) in March. We use these data from 2018 and 2019. Except for the April wave, we included the MHI-5 battery in our CoViD-19 surveys, too.

Figure 2 displays the evolution of the MHI-5 score by gender and calendar month for the working population. At any point in time, men exhibit higher mental health scores than women (individual p-values are highly significant). Prior to the pandemic, values are slightly higher in March compared to November, which is in line with findings from studies on the seasonality of mental health (see,

Figure 2: Evolution of mental health by gender.



(a) Men (b) Women

Note: Lines connect the values of mean mental health during the calender year 2020. Values for 2018 shown in black, for 2019 in red. Vertical bars depict 95-% confidence intervals. Means are estimated on the sample containing the working population.

e.g., Magnusson, 2000, for an overview).

The most salient feature of the graph are the low MHI-5 values in March 2020, which refer to the first two weeks of the initial lockdown. The drop is about 5 points relative to pre-pandemic measures; it is much larger than any other variation there is in this graph. The drop is around a third of a standard deviation. This is a similar order of magnitude as Marcus (2013) finds for unemployed individuals one year after plant closure. In terms of the original questions, it is equivalent to 75% of individuals moving by one step on a question—say, answering "sometimes" instead of "seldom" regarding the statement "I felt very anxious"—and 25% moving two steps (of course, many different ways are possible). We cannot exclude a parallel shift for men and women – subtracting individual values during the pandemic and testing for whether the difference between men and women is different from zero never allows us to exclude the Null hypothesis.

Already in May, mental health was at or above its pre-pandemic levels for the calendar month of March rising more during the summer. Of course, the counterfactual is not fully clear because we lack pre-pandemic measure outside of March or November. Nevertheless, the fact that levels drop again towards November 2020 while point estimates are above pre-pandemic values for the same calendar month leads us to think that average mental health over the summer and autumn is not far from its values in other years. This interpretation is also in line with other studies, which have found mental health to rebound in other countries (e.g., Daly and Robinson, 2021; Daly et al., 2020; Robinson et al., 2022).

2.4 Key explanatory variables

As described in the introduction, we expect that the pandemic-driven infection risk, labor market risk, emotional loneliness, and household structure / childcare arrangements will be key predictors of mental health during the first year of the pandemic. Panel A of Table 2 presents the evolution of measures of risk factors across time and gender. The data at our disposal is unique in that they contain high-frequency data on all these channels along with mental health. We describe the variables we use in turn.

The most direct risk factor of the pandemic for poor mental health is the perceived chance of contracting the virus. We asked respondents about the likelihood of contracting the virus, which they could answer on a scale from 0–100.[2] For November 2019 (pre-Covid) we set the perceived infection risk to zero. The first two rows of Table 2 show the evolution of infection risk on a 0–1 scale. The initial uncertainty surrounding the disease in March 2020 is reflected in a very high perceived chance of contracting the disease. Thereafter, the perceived infection risk tracks infection rates in the Netherlands. With 25-40%, the probabilities seem fairly large throughout, likely reflecting a well-known bias towards 50% (e.g., Wakker, 2010). From May onward, women always perceive a higher chance of being infected with SARS-CoV-2.[3]

A second channel we consider concerns increases in emotional loneliness through the loss of social contacts. In the LISS panel, loneliness is measured with the 6-item De Jong-Gierveld loneliness questionnaire (Gierveld and Tilburg, 2006).[4] Responses on the six items are summed up into a loneliness score ranging from 0 to 12, with higher values indicating a stronger feeling of loneliness. Loneliness was collected pre-Covid and in the Covid waves during April, June, and December. In Table 2 and our main regression specifications, we extrapolate the April values to March and May, and the June values to September. We also run estimations excluding the missing months.

[2] In March 2020, we asked the same question on a 7-point scale, ranging from "no chance" to "certain". We scale this from zero to one as well.

[3] We also asked about diagnosed infection with the Coronavirus. Less than 6% report such a diagnosis at any point in time and 85% of these only do so in December 2020. Due to these (expectedly) small fractions, we refrain from reporting on them in Table 2 or in our main specification below. We provide details on the variable in Online Appendix B while checking robustness to including it among the regressors in Online Appendix C.3.

[4] Participants answer to the six following statements: I experience a void around me, There are plenty of people I can fall back on in case of trouble, I know many people I can fully trust, There are enough people with whom I feel closely connected, I miss people around me, I often feel let down. Answer categories are "yes", "more or less", or "no". For positive items the answers are coded from 0 (yes) to 2 (no). Negative items are coded the other way around.

Table 2: Summary statistics for the key explanatory variables

A. Time-varying measurements							
	Nov ’19	**Mar ’20**	**May ’20**	**June ’20**	**Sept ’20**	**Dec ’20**	**overall**
Perceived CoViD-19 infection risk							
men	0	0.47	0.30	0.25	0.30	0.34	0.28
women	0	0.48	0.33	0.28	0.36	0.39	0.31
Loneliness score							
men	1.86	2.14	2.15	1.96	1.90	1.90	1.99
women	1.86	2.32	2.32	1.94	1.97	1.72	2.04
Labor market outcomes and expectations							
reduction of working hours > 25%							
men	0	0.15	0.23	0.20	0.19	0.15	0.15
women	0	0.23	0.30	0.24	0.21	0.19	0.20
subjective probability of job loss							
men	0.016	0.034	0.033	0.029	0.024	0.026	0.027
women	0.015	0.043	0.024	0.023	0.024	0.017	0.025
Hours worked from home							
men	4.6	17.1	14.7	13.3	10.9	15.2	12.7
women	3.1	11.7	9.5	8.3	6.5	8.9	8.2
Number of observations							
men	832	961	812	795	808	802	5,010
women	859	1,040	842	811	850	792	5,193

B. Household structure / arrangements for extra childcare (measured in March/April 2020)

	child below age 12				**child aged 12-18**	**no child in household**	**single parent**
	jointly	myself	partner	other arrangement			
men	0.10	0.02	0.07	0.04	0.23	0.48	0.05
women	0.09	0.05	0.02	0.05	0.22	0.48	0.09

Note: The information on household structure / caregiver arrangements in Panel B is available for 961 men and 1,040 women.

The second set of rows in Table 2 shows that men and women have similar loneliness scores, just below 1.9 in November 2019 and just above that value in June 2020. For both genders, there is a substantial increase during the first lockdown. In line with Etheridge and Spantig (2022), the resulting level is higher for women (2.3) than for men (2.1).

Third, the pandemic created high uncertainty about individuals labor market prospects. To this end, we construct two measures. First, we use the change in total working hours relative to pre-Covid working hours. This measure is available in all waves and captures direct labor market effects of the pandemic. According to Zimpelmann et al. (2021), this is the most useful measure in the presence of firing restrictions and large-scale economic support programs of the type implemented in the Netherlands.[5]. To measure substantial work disruptions, we construct a dummy variable which takes the value one if the relative reduction in working hours is more than 25%, and is zero otherwise.

To measure labor market uncertainty in the medium run, we use individual's subjective probability of becoming unemployed over the next three months.[6] Unemployment expectations were elicited in all waves but in June 2020, for which we use the numbers from May. As pre-Covid measure, we use job loss expectations from the LISS module "work and schooling". Here, respondents were asked about the subjective probability of losing their job over the next 12 months which is a longer horizon than the three months expectations in the Covid waves. For our main specification, we thus rescale the job-loss expectation for 2019 to reflect the shorter horizon to have a mean of 1.5% instead of 7%.[7]

The last set of rows in Panel A of Table 2 shows the evolution of these two variables. The share of people whose working hours are reduced by at least 25% relative to their respective baseline is 15% for men and 24% for women in March 2020. This share reaches its maximum towards the end of the first lockdown in May (22% and 30% respectively). For men, it remains higher than the initial response throughout the year. By contrast, for women, it falls below the March value in the second half of 2020, but the level remains higher than for men. Subjective job loss probabilities follow a different pattern. For both genders, there is a large peak in March 2020, before gradually falling off over the rest of the year. This temporal variation is substantially more pronounced for women, who have a significantly higher value than men only in March. In December, their perceived probabilities are one percentage point lower than men's. On average, job loss probabilities seem well aligned with actual changes in employment. For instance, the rates of unemployment and of non-participation in the labor force rose by one percentage point each over the course of 2020 (e.g. Meekes et al., 2020; Zimpelmann et al., 2021). Note that the pattern of a shecession found in many countries was not present in the Netherlands (Alon et al., forthcoming).

[5] Non-employment and unemployment rates increased only by about 1 percentage point each in the Netherlands between March and September of 2020, compared to 1.3 and 3.5 percentage points in the U.S., see Bureau of Labor Statistics `https://www.bls.gov/charts/employment-situation/civilian-labor-force-participation-rate.htm` and `https://fred.stlouisfed.org/series/UNRATE`

[6] Such measures have proved to be good predictors of individuals' behaviors with respect to future consequences, such as consumption and savings (see for instance Curtin, 2003; Hendren, 2017; Pettinicchi and Vellekoop, 2019; Stephens Jr., 2004).

[7] In robustness checks, we show that our results are not sensitive to the targeted level by varying it between 0.5% and using the original job-loss expectation measure of 7%. See Online Appendix C.5.

Finally, we measure the extra childcare duties incurred on parents due to the closure of school and daycare facilities. We utilize information on the household structure and care giving arrangements within families. In March 2020, we asked parents of children below the age of 12 how they arrange the *additional* childcare duties during the closure of schools and daycare facilities. Respondents could answer that both respondent and partner take on additional duties (jointly); only the respondent takes on additional duties (myself); only partner takes on additional duties (partner); other all arrangements were made not involving parents (other arrangement).[8]. To take account of other family compositions, we add information on the domestic situation of respondents using the following categories: no children in the household, children older than 12 years, and single parents.

Figure 3 breaks down the evolution of the MHI-5 score by arrangements regarding the extra childcare over the period spanning November 2019 to December 2020.[9] If partners share the additional responsibilities (Figure 3a), there hardly is a drop in the MHI-5 score for men at the beginning of the first lockdown. For women, the drop is somewhat more pronounced. For both genders, the same decline is largest if fathers are responsible for the additional caregiving duties (Figures 3b for men and 3c for women); both recoveries are steep. Our analysis in Section 3 will separate these patterns from other channels.

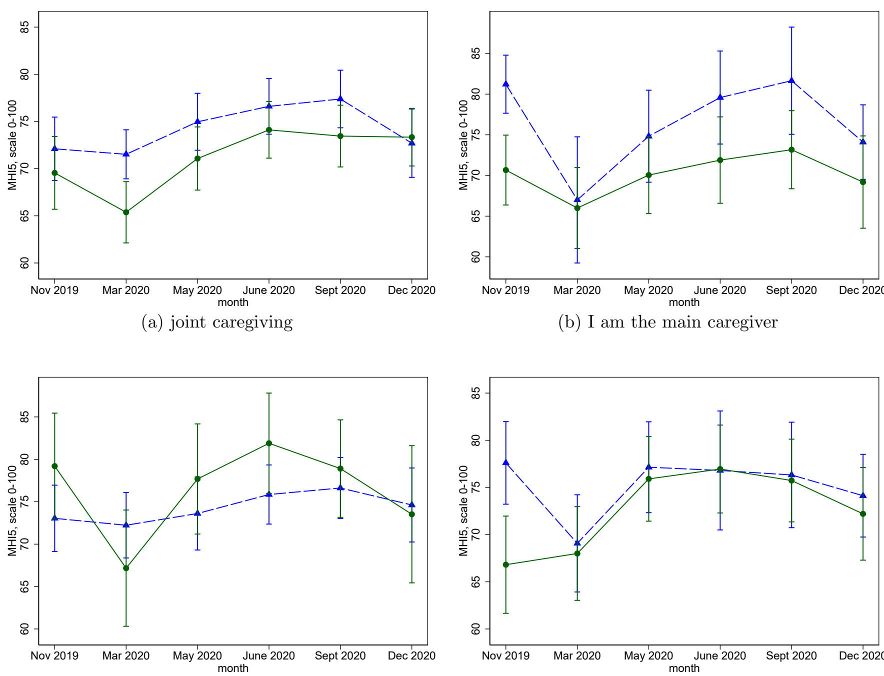
Foreshadowing some of our results, we find that during the initial lockdown, mental health reduces strongly for some groups of parents. In order to dig deeper into the mechanisms, we will do two things. First, we investigate how mental health evolves with the interaction between caregiving arrangements and hours worked from home. The latter are depicted in Panel A of Table 2.

Second, we make use of time use data collected in November 2019 and in April 2020. These two questionnaires were very similar, with some modifications made in the latter survey to adjust for the situation with contact restrictions (see Gaudecker et al., 2020, for the precise wording of the questions). Both questionnaires ask about the number of hours worked at the workplace and about the number of hours worked from home. For parents with young children in the household, the April 2020 questionnaire distinguished between hours worked from home while being responsible for childcare *at the same time* or not. We set this "double burden" variable to zero in November 2019,

[8]The category "other arrangements" summarizes the following non-parental caregivers: older siblings, grandparents, other relatives/friends, ex-partner. It also includes emergency childcare organized by the government for certain professions and that the child stays at home alone.

[9]For reasons of brevity, patterns for types of households other than two parents with small children are relegated to Appendix Figure D.1.

Figure 3: Evolution of mental health by arrangements made for additional childcare duties during school/daycare closures.



(a) joint caregiving

(b) I am the main caregiver

(c) my partner is the main caregiver

(d) external caregiving arrangement

Note: Each panel shows the evolution of mental health separately for men (blue, dashed, triangles) and women (green, solid, circles). Means are estimated on the sample of the working population and conditional on the primary form of care arrangement as stated in March or April 2020. Vertical bars depict 95-% confidence intervals. Similar trajectories for other household structures are shown in Online Appendix D, Figure D.1.

assuming that for the vast majority of hours worked from home, parents had no joint responsibility for childcare.

Table 3: Summary statistics for time use variables

	November '19		**April '20**	
	Men	**Women**	**Men**	**Women**
Hours worked at usual workplace	34.5	26.9	17.4	13.1
Hours worked from home while not being responsible for childcare	3.2	1.9	14.5	10.0
Hours worked from home and responsible for childcare	0.0	0.0	3.4	2.5

Note: Hours refer to a weekly basis. All numbers are averages. Hours working from home while taking care of the children are winsorized above at the 99$^{\text{th}}$ percentile, corresponding to 40 hours per week. The statistics are computed for the sample of 609 men and 602 women who participated in the survey in November 2019 and May 2020 (i.e. the same sample used in Figure 4).

Table 3 shows the sharp increase in hours worked from home between November 2019 and April 2020. Working hours with concurrent responsibility for childare are averaged across all household types. They are, of course, concentrated among parents with young children in the household. In that group, they make up 7.5 and 7.8 hours for men and women, respectively. Note that the April 2020 time use data was collected in the second half of the month and refers to the past week. This week is contained in the four-week recall period for the MHI-5 instrument fielded in May 2020.

3 Individual-level trajectories of mental health

The previous section showed that mental health dropped sharply during the initial phase of the pandemic, rose quickly thereafter, and ended the calendar year at levels typical for that season. In this main part of the paper, we investigate how individual-level trajectories evolved. More precisely, we employ fixed effects methods in order to assess the relative importance of the four channels we scrutinize. In Section 3.1 we present our empirical strategy and describe the results in Section 3.2. Finally, we zoom in on potential mechanisms for how additional childcare duties may have affected mental health. In particular, we investigate the relation between concurrent responsibility for children and working from home and mental health in Section 3.3.

3.1 Empirical strategy

Our main specification is a linear model that exploits the panel dimension of our data to control for time-invariant heterogeneity. To allow the effects of household structure and caregiver arrangements—which we measure in March or April 2020—to vary over time, we add a full set of interactions with survey month fixed effects. We estimate all regressions separately for men and women.

Our model for mental health M_{it} of individual i in survey month t becomes:

$$M_{it} = \delta Inf_{it} + \theta Lonely_{it} + Labor'_{it}\gamma + \sum_{k=1}^{6} \beta_{kt}\mathbb{I}(Caregiver_i = k) \cdot \mathbb{I}_t\eta_t + \alpha_i + \varepsilon_{it}. \tag{1}$$

Inf_{it} is the perceived infection risk; $Lonely_{it}$ represents the loneliness score. $Labor_{it}$ is a vector comprising the indicator for a large reduction in working hours and the subjective risk of job loss. η_t is a survey month fixed effect, α_i is an individual fixed effect, and ε_{it} are unobservables that we assume to be i.i.d. across individuals. $\mathbb{I}(Caregiver_i = k)$ is an indicator taking the value one if individual i falls into one of our seven categories of caregiver arrangements and household types. Recall from section 2.4 that this variable captures the household structure and parents' response to childcare/school/daycare closures, i.e. how parents deal with the *additional* childcare duties. $\mathbb{I}_t$ is an indicator for a particular survey month. The parameters β_{kt} capture the impact of caregiver arrangement k in survey month t on mental health, relative to the respective reference category. We choose joint organization of childcare as the reference category for the caregiver/household type variable and November 2019 as the reference period. We cluster standard errors at the level of the individual. We discuss the results from these specifications in Section 3.2.

In Section 3.3, we investigate the hypothesis whether working from home while being responsible for young children is associated with an increase in the risk of mental health problems. To do so, we interact the number of hours worked from home, $Hrshome_{it}$, with the k categories of childcare arrangements and household types in a first pass:

$$\begin{aligned} M_{it} =& \delta Inf_{it} + \theta Lonely_{it} + Labor'_{it}\gamma + \psi Hrshome_{it} + \sum_{k=1}^{6} \beta_k \mathbb{I}(Caregiver_i = k) \cdot Hrshome_{it} \\ &+ \eta_t + \alpha_i + \varepsilon_{it}. \end{aligned} \tag{2}$$

In this specification, the β_k measure the impact of an additional hour spent working from home for an individual with caregiver arrangement k on mental health. ψ measures the impact of an increase in home office hours for parents who organize childcare jointly during the lockdown. We estimate Equation (2) on two different samples. First, we use the full sample over all survey months. Second, we restrict ourselves to observations until May 2020. This allows us to gauge whether an effect is exclusively due to the period where schools and daycare centers had been closed unexpectedly or whether it persists to a time where they were mostly open but home office hours still high.

Alternatively, we estimate the following regression on the time use data from November 2019 and April 2020:

$$\begin{aligned} M_{it} =& \delta Inf_{it} + \theta Lonely_{it} + Labor'_{it}\gamma + \psi HrshomeChildcare_{it} + \psi_2 HrshomeChildcare^2_{it} \\ &+ \eta_t + \alpha_i + \varepsilon_{it}, \end{aligned} \tag{3}$$

where $HrshomeChildcare_{it}$ measures the hours a respondent works from home while taking care of young children. Note that this specification does not include the household structure / caregiver variable. Instead, information on differences between individuals' risk factors regarding care arrangements comes from the time use data alone.

The interpretation of the coefficients in Equations (1)–(3) is not obvious. Despite the individual fixed effects, they are unlikely to be unbiased estimates of the regressors' causal effects. Reverse causality in particular remains an important concern. For example, a depressive episode or increased anxiety may lead individuals to have a bleak outlook regarding their Covid-19 infection risk or their own future employment. Partners may choose to allocate the additional childcare burden to the individual in better mental health. To address this potential endogeneity issue, we regress the variables measuring each channel of interest (using the earliest observation during CoViD-19) on MHI-5 scores collected in the years 2018 and 2019 and other pre-pandemic factors, relegating the tables with results to the Online Appendix C.1. We employ linear regressions for most variables (Table C.1) and a multinomial logit specification for the choice of caregiver arrangement (Table C.2). The results indicate that for most channels, we are unable to predict the realizations in early 2020 using mental health scores before the pandemic. There are two exceptions. First, individuals in poor mental health are somewhat more likely to experience reductions in working hours down the road.

The most likely explanation here is via socio-ecomonomic status, as low-skilled services jobs were those with the largest hours reductions on average (Zimpelmann et al., 2021). Such explanations points towards the importance of employing fixed effects methods. Second, poor mental health is strongly associated with loneliness down the road, mirroring typical findings (see, for instance, Leigh-Hunt et al., 2017, for an overview). The reverse causality worry is largest for this channel and we will present results below in specifications including and excluding loneliness.

In our view, the overall pattern—i.e., an absence of strong relationships with the exception of loneliness—is not too surprising. For example, the decision on how to allocate the additional caregiving had to be taken extremely fast as schools and daycare centers closed without prior notice. Our data mostly refers to the first week of closures, so there had not been time to re-optimize. Similar arguments can be made for unemployment risk and hours of work. In both cases, the sector of work had very large impacts (e.g. Meekes et al., 2020; Zimpelmann et al., 2021). Nevertheless, this "test" of course does not allow for conclusive evidence from a scientific vantage point. Neither does it preclude the possibility that large drops in mental health may lead to increases in the risk factors. Hence, we will not interpret the coefficients in a causal way. Rather, we think about them as upper bounds on the potential effect sizes. In all examples noted above and indeed all plausible cases we can think of, an effect of each risk factor on mental health would imply the same sign on the coefficient as reverse causality would do.

3.2 Predictors of mental health during the CoViD-19 pandemic

Table 4 presents the estimation results of our main specification for men and women. Because the interaction of time and caregiver / household structure variables leads to a large number of coefficients, we only present those for parents of young children who do not have an external caregiver arrangement in the main text; see Table C.3 in Online Appendix C.2 for the full set of coefficients. Columns (1) and (3) include the specifications with loneliness; Columns (2) and (4) exclude it. In almost all cases, coefficients on other regressors are hardly affected. We thus discuss the results including loneliness, referring to differences caused by its treatment only where it becomes relevant.

We find a significant negative relationship between infections risk and mental health for women. Just opposite, labor market uncertainty is significant only for men. A reduction in working hours

Table 4: Estimated coefficients from fixed effects OLS regressions of mental health on main channels

	Men		Women	
	(1)	(2)	(3)	(4)
prob: becoming infected	-0.90	-0.92	-2.44**	-2.45**
	(0.92)	(0.93)	(1.02)	(1.03)
reduced working hours: yes	-1.26***	-1.27***	-0.74	-0.71
	(0.44)	(0.44)	(0.45)	(0.46)
prob: becoming unemployed	-9.02***	-9.17***	-3.28	-3.41
	(2.10)	(2.06)	(2.13)	(2.17)
loneliness	-0.45***		-0.88***	
	(0.14)		(0.16)	
March 2020 (ref: sharing extra childcare duties)	0.03	-0.21	-0.23	-0.76
	(1.46)	(1.46)	(1.91)	(1.97)
May 2020 (ref: sharing extra childcare duties)	3.09**	2.84**	4.89***	4.37***
	(1.38)	(1.38)	(1.53)	(1.62)
June 2020 (ref: sharing extra childcare duties)	4.41***	4.36***	6.36***	6.25***
	(1.31)	(1.30)	(1.88)	(1.91)
September 2020 (ref: sharing extra childcare duties)	5.21***	5.18***	6.33***	6.17***
	(1.33)	(1.34)	(1.69)	(1.72)
December 2020 (ref: sharing extra childcare duties)	0.10	-0.02	3.33*	3.29*
	(1.49)	(1.47)	(1.92)	(1.96)
extra childcare: myself x March 2020	-12.79***	-12.55***	-2.31	-1.69
	(3.15)	(3.14)	(3.08)	(3.11)
extra childcare: myself x May 2020	-7.51***	-7.26***	-4.90**	-4.26*
	(2.76)	(2.73)	(2.31)	(2.34)
extra childcare: myself x June 2020	-5.80**	-5.88**	-4.58	-4.30
	(2.86)	(2.88)	(3.07)	(3.15)
extra childcare: myself x September 2020	-4.34	-4.51	-4.25	-3.94
	(2.76)	(2.82)	(3.09)	(3.09)
extra childcare: myself x December 2020	-5.56**	-5.51**	-4.71	-4.40
	(2.68)	(2.67)	(3.54)	(3.42)
extra childcare: partner x March 2020	0.22	0.52	-5.66	-5.87
	(2.34)	(2.37)	(3.98)	(4.05)
extra childcare: partner x May 2020	-1.68	-1.42	-4.39	-4.53
	(2.05)	(2.06)	(2.84)	(2.93)
extra childcare: partner x June 2020	-1.49	-1.51	-1.85	-2.25
	(2.25)	(2.27)	(3.50)	(3.59)
extra childcare: partner x September 2020	-1.63	-1.64	-3.75	-4.00
	(1.94)	(1.96)	(3.78)	(3.86)
extra childcare: partner x December 2020	1.90	1.89	-5.40	-5.46
	(2.26)	(2.28)	(3.64)	(3.66)
observations	5,010	5,010	5,193	5,193
number of individuals	1,119	1,119	1,201	1,201

*** p<0.01, ** p<0.05, * p<0.1; standard errors clustered on the individual level; The table presents the estimated coefficients from an OLS regression with individual fixed effects of the MHI-5 score on main channels of the pandemic, see Equation (1). We control for a full set of interactions between survey month and categories of caregiver arrangements and household structure. The reference period is November 2019, the reference category are parents who share the extra childcare that becomes necessary during the closure of school and daycare centers. To economize on space, we do not report all regression coefficients; the full set of results can be found in Online Appendix C.2, Table C.3.

of at least 25% relative to the working hours in the pre-crisis period is associated with a significant reduction in men's mental health score by 1.26 points. An effect of similar magnitude obtains for a fifteen percentage point increase in the probability to lose one's job. For women, the corresponding point estimates are considerably smaller. The difference between genders is statistically significant for perceived unemployment risk (p-value=0.055). It is well-known from earlier work that recessions negatively impact mental health (e.g. Frasquilho et al., 2016; McInerney et al., 2013). Because for most Dutch households, male earnings play a substantially larger role in total household income than female earnings, the gender differences do not come as a surprise.

Neither is it a surprise that reductions in mental health go hand in hand with increases in emotional loneliness. This is the case for both men and women. For men, an increase in the loneliness score by 1 point comes with a decrease in the MHI-5 score of 0.45 points. The same reaction is 0.88 point for women; the difference between genders is significant (p-value=0.045). Given that the scale for loneliness is much larger than that for the other variables (it varies from 0 to 12 whereas other variables are dummies or on a scale from 0 to 1), these are fairly large effect sizes. The gender differences mirror findings for the U.K. reported in Etheridge and Spantig (2022).

The coefficients on the survey month fixed effects show the development of mental health for parents who jointly organized the additional childcare duties caused by the school closures during the first lockdown. After controlling for covariates, the average drop in March is small and insignificant for both genders. For the May-September period, MHI-5 scores are substantially higher compared to November 2019, before falling again. These estimates similar for men and women. By and large, they confirm the patterns outlined in Figure 2.

Relative to this trajectory, parents who were solely responsible for taking on the additional childcare duties experience substantially larger reductions in mental health throughout 2020. The pattern is more pronounced among men, where the average drop in the MHI-5 score between November 2019 and March 2020 is around 12 points. At this early stage of the pandemic, the estimated coefficients for men and women are significantly different on the 5% level (p-value=0.018). The recovery from this shock is slow and significantly worse than for fathers who share caregiver duties with their partners. For women, the initial shock is much smaller; patterns look similar to men from May onward. These coefficients suggest a substantial burden on the mental health of parents in couples where additional childcare is not shared.[10] In December 2020, MHI-5 scores are significantly below their pre-pandemic values for men; point estimates of similar size for women are imprecisely estimated.

As may be expected from Figure 3c, there hardly is a change in men's mental health if their partner has compensated for school and daycare closures. Coefficients are small and insignificant, always working against the hump-shaped pattern. For women whose partner is mainly responsible for additional childcare duties, controlling for covariates cuts the drop upon the onset of the pandemic in Figure 3c by more than half and renders it insignificant. For all coeficients, point estimates are negative and the imprecision inherent in them does not allow us to draw meaningful conclusions.

Summing up, our results show that the patterns from Figure 3 for differences by caregiver arrangement are broadly confirmed in the fixed effects analysis. Moreover, exposure to infection risk and emotional loneliness channel predicts deterioration in mental health among both genders of similar magnitude. For loneliness, the relationship is somewhat stronger among women. By contrast, for men, the pandemic significantly operates through labor market channels. This seems plausible since men are frequently the main breadwinner, implying that the prospect of losing their job may generate more anxiety.

3.3 A double burden of home office and childcare duties?

The results from our main specification have revealed that men who were mainly responsible to handle additional childcare duties experienced the largest initial reduction in mental health. In particular, the drop was larger than for women in the same category. A possible explanation could be that the primary reason for men to exclusively take on the extra childcare is that they can work from home whereas their partners cannot.

In our sample, home office hours for both gender were low before the pandemic. Men worked on average about 4.6 hours from home, women had on average 3.1 home office hours (see Table 2). The pandemic led to a strong increase in hours worked from home. In March 2020, men on average worked 17.1 hours from home and women worked 11.7 hours from home. Over the summer, the

[10] Unfortunately, the small sample size prevents us to investigate further to what extent these choices are deliberate or heavily influenced by constraints such as work schedules of essential workers early in the pandemic.

numbers decreased to an average of 10.9 hours for men and 6.5 hours for women, before the rose again in December 2020. Throughout the pandemic the absolute number of hours worked from home was constantly about 40% higher for men than for women. The difference in hours worked from home is more uneven when taking account of caregiver duties. Among families with fathers being the main caregiver, they work 23 hours from home during the March-December 2020 period whereas mothers' home office hours are just below 4. At the same time, there are no reductions in working hours for parents of underage children relative to the remaining population (Holler et al., 2021). Hence, many men in this group will be faced with the task of working and taking care of their children at the same time.

Table 5 reports the results of estimating Equation (2), i.e., fixed effects regressions including an interaction of home office hours and the extra caregiver /household structure variable. Panel A shows the most important results for the full set of time periods; Panel B focuses on the comparison between November 2019 and the first lockdown period, which included closed schools and daycare.

Among the reference group—couples sharing the extra childcare burden—hours worked from home do not bear a significant relation with mental health. For both genders, the coefficients are close to zero and precisely estimated. When fathers take over the main caregiver responsibility, an extra hour worked from home is associated with a reduction of 0.2 points in the MHI-5 score. On average, these fathers work about 20 hours from home, so their mental health is reduced by roughly 4 more points on average compared to similar men in couples who share the additional responsibilities. The point estimates for women are smaller and imprecisely estimated. Neither is it significantly different from zero nor from those for men (p-value=0.47).

For fathers whose partner takes on the main responsibility for additional childcare, we cannot detect a differential relation between hours worked from home and mental health. By contrast, women experience a large and significant decrease in mental health as home office hours increase. This decrease in mental health is significantly larger for women than for men (p-value=0.004 in Panel A and p-value=0.003 in Panel B). This effect is substantially larger when concentrating on the initial lockdown period in Panel B. Together with the findings in Figure 3c and Table 4, this coefficient suggests that women are particularly at risk for developing mental health problems in situations where partners are supposed to take care of the children while they work from home.

Finally, we take another look from a different angle using a direct measure for hours worked

Table 5: Estimated coefficients from fixed effects OLS regressions of mental health on hours worked from home by arrangement for extra childcare duties

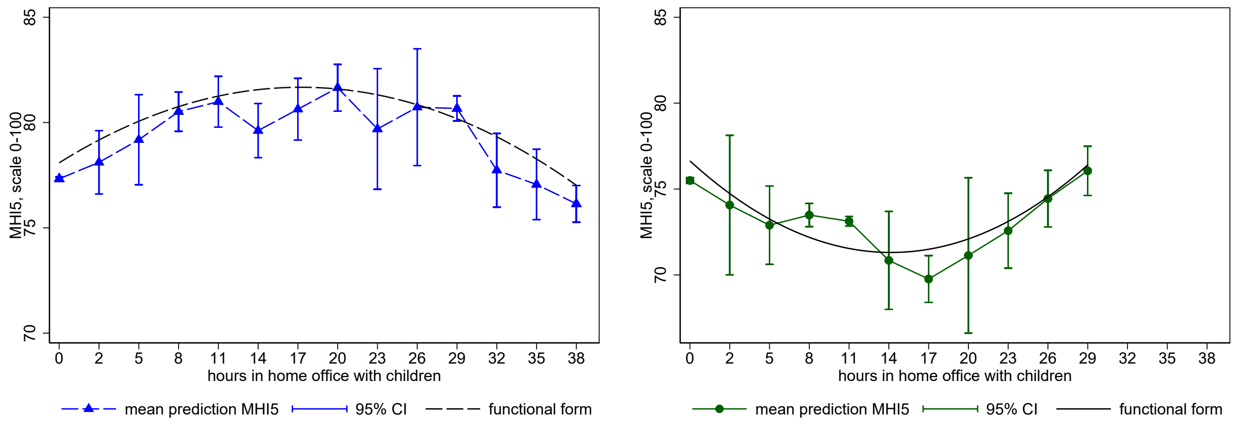
	Men		Women	
	(1)	(2)	(3)	(4)
A. all periods				
hours worked from home (reference: sharing extra childcare duties)	-0.02	-0.02	0.01	0.00
	(0.03)	(0.03)	(0.05)	(0.05)
extra childcare: myself x hours worked from home	-0.19***	-0.19***	-0.09	-0.08
	(0.06)	(0.06)	(0.12)	(0.12)
extra childcare: partner x hours worked from home	0.00	0.00	-0.41***	-0.43***
	(0.06)	(0.06)	(0.13)	(0.14)
observations	5,010	5,010	5,193	5,193
number of individuals	1,119	1,119	1,201	1,201
loneliness among covariates	yes	no	yes	no
B. periods before the pandemic and with closed schools/childcare facilities				
hours worked from home (reference: sharing extra childcare duties)	0.03	0.02	-0.06	-0.08
	(0.05)	(0.05)	(0.09)	(0.09)
extra childcare: myself x hours worked from home	-0.23**	-0.22**	-0.08	-0.03
	(0.10)	(0.10)	(0.15)	(0.15)
extra childcare: partner x hours worked from home	-0.08	-0.06	-0.77***	-0.77***
	(0.10)	(0.10)	(0.21)	(0.22)
observations	2,605	2,605	2,741	2,741
number of individuals	1,099	1,099	1,179	1,179
loneliness among covariates	yes	no	yes	no

*** p<0.01, ** p<0.05, * p<0.1; standard errors clustered on the individual level; the table presents the estimated coefficients from OLS regressions with individual fixed effects of the MHI-5 score on interactions between home office hours and extra care duties for men and women, see Equation (2). The reference period is November 2019, the reference category are parents who share the extra childcare that becomes necessary during the closure of school and daycare centers. Panel A shows the regression results for the full sample period from November 2019 to December 2020. Panel B shows the results from corresponding fixed effects OLS regressions when we restrict our sample to November 2019 and the first Covid wave from March to May 2020. The table only presents the coefficients on home office working hours and interactions between home office working hours and the set of childcare arrangements for children under 12 years of age. The full set of interactions, infection risk, labor market channel and social interaction channel is shown in the Online Appendix C.2, Table C.4.

from home while being responsible for childcare at the same time. This measure is included in the time use survey in April 2020. The survey similar to the one from November 2019 but adapted to the lockdown situation (see Gaudecker et al., 2020, for the precise wording of the questions). Time use refers to the past week; this week falls into the four-week assessment period for mental health in the November and May questionnaires, respectively.

Figure 4 visualizes the key results from estimating Equation (3), i.e., a fixed effects regression of mental health on a quadratic in hours worked from home while being responsible for childcare at the same time. The regression specification also includes the remaining three channels and survey month fixed effects. We plot up to the 99^{th} percentile of the distribution of hours worked from home with kids present, which is 40 hours for men and 30 hours for women. For men, we

Figure 4: Predicted mental health score by hours worked from home while simultaneously taking care of children.



(a) Men (b) Women

Note: The figure plots predicted mental health against hours in home office while taking care of children. Predicted values are obtained from an OLS regression with individual-specific fixed effects of mental health on a quadratic in hours worked from home while being responsible for children at the same time. The estimation is based on a sample of 609 men and 602 women who participated in the survey in November 2019 and May 2020. We control for measures of labor market risk, infection risk, and social interaction channels, and survey month fixed effects. The estimated coefficients from the quadratic specification can be found in Online Appendix C.2, Table C.5. We use the average of fixed effects to adjust the level of mental health based on the quadratic function to those in the data. The predicted values use bins of three hours.

find a hump-shaped relationship between mental health and home office hours, which reaches its maximum around 18 and its minimum in the right tail of the distribution. For women the pattern looks opposite, suggesting that women who work around 15 hours from home and take care of their children at the same time have the lowest mental health score. One difference between genders is the respective peak / trough of mental health around 15-18 hours which makes up 60-70% of these women's working time, whereas it is only around 40% for these men (see Figure D.2 in Online Appendix D).

Importantly, there are no systematic differences in November 2019—neither in mental health nor working hours—along the distribution of hours worked from home with kids. In particular, in November 2019, mental health is about the same whether or not somebody reports positive home office hours with children in April 2020. The patterns thus do not seem to be driven by selection or regression to the mean. Furthermore, total working hours in April do not have a clear relationship with the amount of home office hours with children. To be precise, holding the ability to work from home constant by conditioning on positive hours worked from home, there is no difference between parents who mind their children at the same time and workers who never do so.[11]

[11] Among all mothers working from home, those who also take care of their children work on average 31.3 hours in total. Women who do not mind their children at the same time work 31.2 hours. Fathers in home office work on

3.4 Robustness

Our main results are robust to various alternative ways to specify our regressions. The tables in the main text have already revealed that it is immaterial whether we include loneliness among the regressors or not. In March 2020, the perceived infection risk was asked on a 7-point scale with an additional option that the respondent was already contracted with the virus. In our main specifications, we assigned this outcome to 1, i.e., the same value as "certain". However, prior infection with Covid-19 may imply a change in views on the risks of re-infection or have direct effects on mental health. We thus re-estimate Equation (1), adding a separate dummy variable for those who have already been infected by the virus. The results can be found in Table C.6 in Online Appendix C.3. None of the point estimates changes in a meaningful way. Moreover, a Covid infection does not have a significant impact on mental health of men and women.

In a second robustness check, we estimate Equations (1) and (2) on a balanced sample. Keeping only respondents who participated in all waves reduces the sample to 480 men and 455 women. Table C.7, Online Appendix C.4, presents the estimation results corresponding to Table 4. The only meaningful change in the coefficient estimates is for the perceived infection risk. Here, the estimated coefficient for men becomes large and significant while the one for women is now small and insignificant. All other coefficient estimates somewhat change magnitude but otherwise show the same pattern as for our main specification. Table C.8 presents the balanced-sample results for Equation (2). The estimated coefficients are very similar to those in Table 5, suggesting that our main results are robust to changes in the sample composition.

Next, we investigate whether our results are sensitive to the targeted level at 1.5% of job-loss expectations in 2019. To this end, we vary the target level between 0.5% and original measure of 7%. The results are presented in Online Appendix C.5. The estimated coefficient for the job-loss probability is smaller when using the original job-loss expectation measure, but remain highly significant for men. We do not find any major changes for other target levels. Neither are other coefficients affected by these alternative calibrations.

Finally, we use a binary indicator for whether a person is at risk of developing mental health

average 40.1 hours when not taking care of their children at the same time compared to 36.8 hours when also taking care of children. These differences in working hours are about the same for mothers (30.1 vs 31.2) and fathers (41.8 vs 36.8) in November 2019.

problems as the dependent variable and re-estimate Equations (1) and (2). The binary indicator takes the value 1 if a respondent's MHI-5 score is below 60, and is zero otherwise. Thus, positive coefficients refer to an increase in probability of having mental health problems. Tables C.15 and C.16 in Online Appendix C.6 present the results. Compared to the main estimates in Tables 4 and 5, some estimated coefficients are no longer significant. This is to be expected as we are reducing the information in the regression. The direction of relevant coefficients is mostly the same.

4 Discussion and conclusions

We have analyzed how changes in the mental health of a representative sample of the Dutch working population evolved from before the CoViD-19 pandemic through its first year. Upon the onset of the first lockdown, amidst a period of high uncertainty in many dimensions, mental health dropped very sharply. It recovered over the summer before dropping slightly again, so that December 2020 values are comparable to those from November 2019. Investigating the joint evolution with several potential mediators identified in the literature—household structure and arrangements for taking care of children during the period of school closures, SARS-CoV-2 infection risk, employment prospects, and lack of social interactions—we document substantial heterogeneity.

Mental health falls in perceived infection risk, maybe more so for women. Increases in emotional loneliness are associated with drops in mental health for both genders, but more so for women. On the other hand, the effects of labor market risk are substantially more pronounced for men, which is consistent with them contributing the larger share of income in most families.

The hump-shape of the MHI-5 evolution over the March-December period is more pronounced for parents. The onset of the spring lockdown was a particularly stressful period for them. They had to cope with closed schools and daycare facilities from one day to the next while managing their usual work at the same time. We do not find clear gender effects and in fact, some of the largest drops are found for fathers when they are solely responsible for the additional childcare.

In apparent contrast to this, much of the international literature on mental health problems during the first year of the pandemic has documented that women were more at risk than men. The contrast is only apparent because our results are conditional on individual fixed effects. Also in our data, women have lower *levels* of MHI-5 scores than men at any point in time. Running a regression as Zamarro and Prados (2021, Equation (5) / Table 7), we find similar effects for late March as they do for early April (see Table C.17 in the Online Appendix): Women are significantly more likely to be at risk for mental health problems and the effect is more pronounced for mothers of school-aged children. Interpreted in conjunction with our remaining results, much of this can be attributed to baseline differences rather than differential responses in the Dutch data.

Our study is closest to Etheridge and Spantig (2022) who use the Understanding Society Panel in the U.K. to to document changes in mental health. They find no baseline differences between genders, but much larger drops for women. Decomposing the differences, they trace them back to larger influences of loneliness and child care responsibilities for females. The direction of these effects is consistent with our findings, but the quantitative implications are very different. We speculate that this likely has to do with a very high share of part-time work among women (more than 60% worked less than 30 hours per week in 2017, see OECD, 2018) and very flexible work arrangements, which are mandated by a 2016 law.

In general, our results paint a nuanced picture of the effects of the pandemic in two-parent families, which depend on the degree the extra burden during school and daycare closures is shared between partners and on the fraction of working time that is performed from home while simultaneously being responsible for children. Taken together, our results are consistent with literature showing large but transitory impacts of negative aggregate shocks on mental well-being. For example, Deaton (2012) finds a large impact of the Great Recession in late 2008 and early 2009. These values subsequently recovered despite the fact that unemployment remained high. During the CoViD-19 pandemic in the Netherlands, mental health indicators substantially improved for parents after the period when schools were closed. Despite an imminent second lockdown in December, mental health in the working population was similar to what it was before the pandemic. Our results are perhaps best explained by short-run anxiety associated with a novel and negative situation characterized by uncertainty and quick subsequent adoption.

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