The London School of Economics and Political Science

Working Hours, Childcare Support, Wage Inequality and Windfall Gains

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Declaration

I certify that the thesis I have presented for examination for the MPhil/PhD degree of the London School of Economics and Political Science is solely my own work other than where I have clearly indicated that it is the work of others (in which case the extent of any work carried out jointly by me and any other person is clearly identified in it).

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Abstract

This work analyses working hours, childcare support, wage inequality and windfall gains.

In Chapter I, I test whether family-support policies play a role in explaining variation in working hours across countries. I analyse childcare subsidies and family cash benefits and I distinguish between people with children and people without children. Childcare subsidies should increase working hours in the economy and these effects should differ between parents and nonparents. I test this using household data for a set of European countries and the US. Empirical analysis, however, does not support the family-policy explanation. The effects of the policies on working hours are weak and insignificant. Furthermore, I do not find evidence for the expected differences between parents and nonparents. I conclude that family policies are not helpful in explaining the variation in working hours across countries.

In Chapter II, I argue that rising inequality in offered wages lowers average working hours. If the labour supply is concave in wages, the aggregate effect of the decrease in working hours of low-paid workers is greater than the increase in working hours of high-paid workers. Furthermore, due to low market opportunities, some of the low-paid workers may leave the labour force and become inactive. Using the CPS-MORG data for prime-age men I find evidence in support of this explanation. After controlling for the average wage, wage inequality has a negative effect on the labour supply.

In Chapter III, I investigate whether workers adjust hours of work in response to windfall gains using data from the European Household Panel. The results suggest that unexpected variation in income has a small negative effect on working hours. Furthermore, the empirical findings show that the impact of windfall gains is more important for young and old individuals, is most negative for married individuals with young children, but can be positive for single individuals at the age of 40.

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Doing a PhD is a unique experience. I have met not a single person who can describe their experience with simple words or straightforward emotions. I am nevertheless confident that after these long years of bewildered sentiment, failure and success, pride and shame, student craziness and academic bluntness, I am emerging a better person.

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General Introduction

This work comprises of three papers of which the main topic is working hours. It is important to study working hours for several reasons. In many settings working hours represent an argument in the utility function of individuals, hence they directly affect welfare. Working hours are also a direct demonstration of the labour supply, and in this way they are central to the workings of the labour market and the whole economy. In this work I analyse working hours in relation to childcare support, wage inequality and windfall gains.

In Chapter I, I test whether family-support policies play a role in explaining variation in working hours across countries. There are large differences across countries in the actual hours worked per person. According to OECD data, working hours per person are on average about 40% higher in the US than in Belgium, Germany or France. Even within Europe itself, differences are substantial; an average person in the UK or Scandinavia works almost 30% more than the average person in continental Europe. There is large body of research that has tried to explain this variation. Prescott (2004) argues that such a big US-EU gap in working hours can be explained almost completely by the higher taxes prevalent in Europe. However, while the tax story is plausible when comparing the US with European countries, it cannot accommodate the variation in hours within Europe. More specifically, it fails to explain the fact that in Scandinavia, despite facing higher taxes, people on average work more than in continental Europe.

In response to this Ragan (2005) and Rogerson (2007) propose an extension of the simple tax story. They argue that it is not only important how big the tax burden is, but also how the government uses its tax proceeds. If taxes are used to support disabled people who cannot work, for example, this would increase the differences in hours worked. On the other hand, if taxes are used to subsidise day care for children of parents who work, this would dampen the effect of taxes on work-hours differences. They expand the model to allow for the service sector that produces childcare and other home care services. By means of calibration and simulation they show that accounting for stylised differences in public policies, including the public provision for household services (i.e. childcare), they can much better fit the data on working hours across countries.

The purpose of Chapter I is to identify the role of family policies in explaining working-hours variation across countries. Prescott (2004), Ragan (2005) and Rogerson (2007) focus on simulations of calibrated macroeconomic models with a representative agent. However, there is an underlying logic behind the aggregate effects that they consider. For example, higher childcare subsidies increase aggregate working hours. But the reason for this is that with higher childcare support parents are stimulated to work more, therefore they are the ones for which we should observe large increases in working hours. In my approach I thus take one step back; I ask who and how is affected by the family policies, and whether evidence of this can be found in the data rather than by means of simulation and calibration.

First, I focus on comparisons of two groups, people with children (parents) and people without children (nonparents). Intuitively, family policies should have most of their effects concentrated on the recipients: people with (young) children. Second, I focus my attention on two types of policies: provision of childcare services by the government (childcare subsidies) and direct cash transfers to families (cash benefits). I develop a simple labour supply model with taxes, childcare subsidies and family cash transfers and with two types of households: parents and nonparents. The two types differ in their utility towards childcare and in their budget constraints due to differences in policy receipts. I show that, holding taxes fixed, one would expect a positive effect of childcare subsidies and a negative effect of cash benefits on aggregate hours worked in the market. Furthermore, effects of each policy on working hours of parents should be stronger than effects on working hours of nonparents. Another interesting prediction from the model concerns the effects of family policies on the time spent in childcare at home. Childcare subsidies are expected to decrease aggregate childcare at home, while family cash benefits should increase it; the effects are again expected to be stronger for parents.

If the family-policy explanation of working-hours variation across countries is correct, then the above listed predictions should show up empirically. I obtain data on family support policies across countries from the OECD Social Expenditure database. In order to distinguish between people with children and people without children I use individual level data from the European Household Panel (ECHP) and the US CPS, available for the years 1998 - 2001. Due to the short time period of the available data and the low variation in policy variables over time, I mostly exploit cross-country variation in my analysis.

The results, however do not support the idea that family support policies can help explain differences in working hours across countries. In the preferred specifications with working hours as a dependent variable, the estimated effects for the parents are statistically insignificant and close to zero. Whereas in countries with higher childcare subsidies participation in the labour force is higher, working hours conditional on working are lower, bringing the aggregate effect of policies to zero. In regressions with time spent in childcare at home on the left hand side, the effects of policy variables contradict the theory. An analysis performed on the sample of people older than 55 years, for which family policy should not matter much, gives rise to evidence that in countries with more generous family policies, time spent in childcare of older people is shorter. One possible explanation for this is that public childcare support actually substitutes for the childcare by older people, rather than for the time spent in childcare of the prime-age group. As a majority of older people do not participate in the labour market, family support thus cannot have a strong impact on working hours.

There could be important cultural differences across countries that my analysis does not capture. Therefore I include into the analysis also two measures of "culture": average number of adults in the household and the incidence of part-time work in the country. After including these controls, the aggregate effects of childcare subsidies on working hours become positive and significant; however the effects on parents are still weaker than the effects on nonparents. Furthermore, the results from the analysis on the time in childcare spent at home still contradict the theory.

In Chapter II, I argue that rising inequality in offered wages lowers average working hours. In my work I conceptually follow the "conventional" labour supply literature such as Juhn, Murphy and Topel (1991), Juhn (1992) and Welch (1997). In this type of setting an individual's labour supply simply traces the wage that an individual (potentially) earns in the labour market. Assuming a positive effect of wages on working hours, if the offered wage of a worker rises, the labour supply of a worker will rise. On the other hand, if the offered wage falls, the worker will decrease her work effort or may drop out of the labour force altogether. If in addition the labour supply function is stable over time, then movements in a worker's wage will closely determine movements in his or her labour supply over time.

According to such a view of the labour market, it is obvious that changes in the wage inequality should have a direct impact on labour supply. I argue that a mean preserving spread in the wage distribution causes average working hours in the economy to fall. The reasoning behind this is quite simple. Suppose that the labour supply function is concave in offered wages. When variation in wages increases, keeping the mean wage constant, high-paid workers are getting paid more whereas low-paid workers are getting paid less. Due to the concavity of the labour supply, effects on the labour supply at the bottom of the wage distribution will be stronger than at the top of the wage distribution. Therefore, while high-paid workers increase the number of hours supplied to the market by little, low-paid workers decrease their hours supplied to the market relatively more. Moreover, for some workers wages are reduced below their reservation wage and they decide to drop out of the labour force and work zero hours. In practice, active workers usually work a certain amount of hours that is not close to zero. Hence, a decision to leave the labour force usually represents a discreet jump and reduces working hours substantially. By this argument, therefore, higher variation in wages causes working hours in the economy to fall.

For the results of Chapter II I rely heavily on the concavity of the labour supply function. In fact, the concavity of the labour supply is something that has been around in the literature, at least implicitly, for some time. Much of the established labour-supply literature assumes a concave labour supply, as reported also in Blundell and MaCurdy (1999). In some cases, the concavity of labour supply was reported also explicitly: Juhn et al. (1991) and Juhn and Murphy (1997) find that with rising wages labour supply elasticity falls. Furthermore, in the empirical section of Chapter II, I find support for the concavity of labour supply with my data.

I report empirical evidence in support of the claim that greater wage inequality decreases average hours of work in the economy. I do this using the NBER extracts of the CPS Annual Earnings File (also known as the Merged Outgoing Rotation Groups) for prime-age men for the period 1979-2008. I find evidence that after controlling for the average wage, wage inequality has indeed a negative and significant effect on the labour supply. This result is robust to various specifications. According to my results, the increase in wage inequality equivalent to the increase in the US over the 1979-2008 period would cause an approximately 3% decrease in average hours worked in the economy. In the same period average working hours in the economy fell by about 8%. From this it follows that rising inequality can potentially explain 30-40% of the decrease in working hours over time.

The findings of Chapter II are important for several reasons. The chapter deals with two central themes in labour economics – labour supply and inequality. To my best knowledge this is the first paper that explicitly relates the conventional static view of labour supply, where labour supply is determined by a worker's own offered wage, with the wage inequality. The results are also interesting from the point of view of public policies that affect wage inequality.

In Chapter III, my co-author and I investigate whether workers adjust hours of work in response to unanticipated windfall gains. According to the life-cycle model, a relaxation or tightening of the consumer's intertemporal budget constraint can lead both to changes in consumption and to changes in labour supply. Windfall gains represent an unanticipated increase in non-earned income and, by reducing an agent's marginal utility of wealth, they therefore reduce her incentive to work.

We analyze the linkages between windfall gains and working hours using data from the European Community Household Panel from 1994 to 2001. We show that an unanticipated rise in wealth reduces working hours in accordance with the life-cycle model, although the effect is, in general, small. The impact of windfall gains is stronger at the external margin, that is, individuals adjust their labour supply primarily by entering or dropping out of the labour force, rather than by reducing their work hours conditional on working.

We also examine whether "size matters" with respect to the effects of windfalls on working hours. We assess how households respond to small, medium or large windfall gains. We find that the effects become stronger as the size of the windfalls increases. In particular, men receiving a large windfall on average reduce labour supply by 1.3 hours per week, a 3.4% reduction in working hours.

Finally, analysing the effects of windfall in relation to various personal characteristics, we find that: (i) at younger and older ages, the effect of windfall gains on labour supply is the most negative; (ii) for married people and people with young children the windfall gain leads to a stronger decrease in working hours and (iii) for single individuals at the age of around 40 the effect can be positive. A potential explanation for the latter empirical finding is in the effect of windfall gains on reducing liquidity constraints in capital markets. By doing so windfall gains may encourage people to become self-employed and increase their working hours, as suggested by Lindh and Ohlsson (1996) and Taylor (2001) and also confirmed in our data.

Chapter III contributes to the literature in following ways. This is the first paper that analyses the effects of windfall gains on working hours using European data with more than one country included. Henley (2004) was the first paper done on European data, but it analyzes the effects of capital gains on labour supply using data for Britain only. Furthermore, because we include 15 countries in our analysis, the sample of people for which we observe windfall gains is large, offering a further empirical advantage to our approach. With the panel data set we additionally observe a rich set of personal characteristics of individuals. This gives us an opportunity to better understand the ways that participation and working-hours decisions vary across different types of individuals.

Chapter I

Can Family-Support Policies Help Explain Differences in Working Hours across Countries?

Urban Sila

Abstract: It has been suggested in the literature that taxes and subsidies play an important role in explaining the differences in aggregate working hours across countries. In this chapter, using individual level data for a set of European countries and the US, I test whether public programmes for family support play a role in explaining this variation. Extending a simple macroeconomic model, I analyse two types of policies: childcare subsidies and family cash benefits. I distinguish between people with children and people without children. Childcare subsidies should increase working hours in the economy and these effects should differ between people with children and people without children. Public support to families is also expected to decrease the amount of time people spend in childcare at home. The empirical analysis, however, does not support the family-policy explanation. The effects of the policies on working hours are weak and insignificant. In regressions with time spent caring for children as a dependent variable, the estimates of the effects contradict the predictions of the theory. Furthermore, I do not find evidence for the expected differences in effects between parents and nonparents. I conclude that family policies are not helpful in explaining the variation in working hours across countries.

1 Introduction

There are large differences across countries in the actual hours worked per person. According to OECD data, working hours per person are on average about 40% higher in the US than in Belgium, Germany or France. Even within Europe itself, differences are substantial; an average person in the UK or Scandinavia works almost 30% more than the average person in continental Europe. There is large body of research that has tried to explain this variation. One branch of the literature explains the "hours divide" with differences in tax rates. Prescott (2004) relies on calibration techniques and argues that such a big US-EU gap in working hours can be explained almost completely by the higher taxes prevalent in Europe. Olovsson (2004) similarly finds that much of the gap between the US and Sweden can be explained by differences in tax rates. Estimates in Davis and Henrekson (2004) provide evidence that taxes significantly reduce the hours worked. However, while the tax story is plausible when comparing the US with European countries, it cannot accommodate the variation in hours *within* Europe. More specifically, it fails to explain the fact that in Scandinavia, despite facing higher taxes, people on average work more than in continental Europe.

In response to this, Ragan (2005) and Rogerson (2007) propose an extension of the simple tax story. They argue that it is not only important how big the tax burden is, but also how the government uses its tax proceeds. While in the model of Prescott (2004) all taxes are returned into the economy as a lump sum subsidy, in reality policies are more complicated in ways that are important for the working-hours variation. If taxes are used to support disabled people who cannot work, for example, this would increase the differences in hours worked. On the other hand, if taxes are used to subsidise day care for children of parents who work, this would dampen the effect of taxes on work-hours differences. This extension can thus be used to better explain the variation in working hours with respect to Scandinavia, where high taxes coincide with generous system of public provision of family services[1]. Ragan (2005) and Rogerson (2007) expand the model to allow for a service sector that produces childcare and other home care services. By means of calibration and simulation, they show that accounting for stylised differences in public policies, including the public provision for household services (i.e. childcare), they can much better fit the data on working hours across countries[2].

[1] Rosen (1996) discusses Swedish case.

[2] A similar story is offered in Ngai and Pissarides (2008). They examine the implications of tax and subsidy policies for employment in three stylized worlds of welfare: a world with low taxation and low regulation of market work (i.e. the UK and the US), a world with higher taxation and regulation (i.e. France and Italy), and a

The purpose of this chapter is to identify the role of family policies in explaining working-hours variation across countries. In this way, I am contributing to the macroeconomic literature. Prescott (2004), Ragan (2005) and Rogerson (2007) focus on simulations of calibrated macroeconomic models with a representative agent, but such aggregate effects are hard to identify empirically with the limited available data. However, there is an underlying logic behind the aggregate effects that they consider. For example, more generous childcare policies increase aggregate working hours. But the reason for this is that with higher childcare support parents are stimulated to work more, therefore they are the ones for whom we should observe large increases in working hours. As such differences may be easier to capture empirically, I step away from the simplest macroeconomic model with a representative agent. At the same time, however, it is important to remember that throughout the chapter I address a macroeconomic question and that my interest lies in explaining the variation in *aggregate* working hours.

First, I focus on comparisons of two groups, people with children (*parents*) and people without children (*nonparents*). Intuitively, family policies should have most of their effects concentrated on the recipients: people with (young) children. Second, I focus my attention on two types of policies: provision of childcare services by the government (*childcare subsidies*) and direct cash transfers to families (*cash benefits*). I develop a simple general equilibrium model that focuses on labour supply and includes taxes, childcare subsidies and family cash transfers as well as two types of households: parents and nonparents. The two types differ in their utility towards childcare and in their budget constraints due to differences in policy receipts. I show that, holding taxes fixed, one would expect a positive effect of childcare subsidies and a negative effect of cash benefits on aggregate hours worked in the market. Furthermore, the effects of each policy on the working hours of parents should be stronger than the effects on the working hours of nonparents. Another interesting prediction from the model concerns the effects of family policies on the time spent in childcare at home. Childcare subsidies are expected to decrease aggregate childcare at home, while family cash benefits should increase it; the effects are again expected to be stronger for parents.

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world with high taxation and regulation, but also with high public support of market services connected with childcare (i.e. Sweden). They claim that taxes and subsidies have different effects on market activity in different sectors of the economy, depending on how close a substitute the product of a particular sector is to home production. They argue that in this way they can explain quite well the differences in employment rates in different sectors between the three worlds of welfare.

If the family-policy explanation of working-hours variation across countries put forward by Prescott (2004), Ragan (2005) and Rogerson (2007) is correct, then the above listed predictions should show up empirically. I obtain data on family support policies across countries from the OECD Social Expenditure database. In order to distinguish between people with children and people without children I use individual level data from the European Household Panel (ECHP) and the US CPS, available for the years 1998 - 2001. Due to the short time period of available data and the low variation in policy variables over time, I mostly exploit cross-country variation in my analysis.

The results do not support the idea that family support policies can help explain differences in working hours across countries. In preferred specifications with working hours as the dependent variable, the estimated effects for the parents are statistically insignificant and close to zero. Whereas in countries with higher childcare subsidies participation in the labour force is higher, working hours *conditional on working* are lower, bringing the aggregate effect of policies to zero. In regressions with time spent in childcare at home on the left hand side, the effects of policy variables contradict the theory. An analysis performed on the sample of people older than 55 years, for which family policy should not matter much, gives rise to evidence that in countries with stronger family policies, older people spend more time in childcare. One possible explanation for this is that public childcare support actually substitutes for the childcare by older people, rather than for the time spent in childcare of the prime-age group. As the majority of older people do not participate in the labour market, family support thus cannot have a strong impact on working hours.

There could be important cultural differences across countries that my analysis does not capture. I therefore also include two measures of "culture" in the analysis: the average number of adults in the household and the incidence of part-time work in the country. After controlling for these variables, the aggregate effects of childcare subsidies on working hours become positive and significant; however, the effects on parents are still weaker than the effects on nonparents. Furthermore, the results from the analysis on the time in childcare spent at home still contradict the theory.

Before turning to the paper, it should be mentioned that the story of taxes and family-policy is not the only story that has been put forward as an explanation of cross-country differences in working hours; on the contrary, the literature is quite large. Faggio & Nickell (2007) have a good statistical survey of the differences in working time across countries and they review and discuss various explanations. Alesina et al. (2005) rely on the Blanchard & Wolfers

(2003) explanation that different institutional settings result in different reactions to adverse shocks; i.e. in Europe unionisation and labour market regulation increased sharply after the shocks of the 1970s and 1980s. The shocks resulted in the so called "work-sharing" policies, which can be used to explain the low hours worked in Europe. Bell & Freeman (1994, 2001), Freeman & Schettkat (2001a) and Schettkat (2003) relate working-hours variation to earnings inequality, where earnings inequality increases work effort due to higher incentives. Bowles & Park (2005) also focus on inequality, but claim that the positive relationship stems from the fact that social comparisons work upwards and hence the poor want to imitate the rich in their consumption. Freeman & Schettkat (2001b, 2005) explain the differences between the US and Europe by the marketisation hypothesis, stating that households in Europe produce more goods at home than the US.

Another branch of literature finds explanations for differences in working hours across countries and over time in factors such as differences in technologies, growth rates and structural transformation (Rogerson (2008, 2006), Pissarides (2007)). Related literature focuses on explaining the trends in the US data (McGrattan & Rogerson (2004), Greenwood & Vandenbroucke (2005), Aguiar & Hurst (2006), Francis & Ramey (2006), Ngai & Pissarides (2007)). Attanasio et al. (2008) try to explain changes in female labour force participation in the US within the framework of a life-cycle model. They argue that female participation has increased for recent cohorts due to reduced cost of childcare and rising wages.

Finally, there is also an extensive literature seeking to assess the microeconomic impacts of family-support policies (free/subsidized childcare) on the labour market outcomes of families, mainly mothers with pre-school children. However, empirical estimates of the effects of these policies on the labour supply of mothers have a wide range, from large to insignificant. For some examples of this literature refer to Baker et al. (2008), Lefebvre and Merrigan (2008), Gelbach (2002), Blau (2000) or Anderson & Levine (1999).

The structure of this chapter is as follows. In Section 2 I introduce a simple theoretical model and discuss its implications. Section 3 and Section 4 show the variation in working hours and the variation in public family policies across countries, respectively. In Section 5 I undertake an empirical analysis of the effects of family policies on working hours, distinguishing between people with children and people without children. In Section 6 I analyze the effects of family policies on the time spent in childcare at home. In Section 7 I add into the analysis two country-level measures of culture and Section 8 concludes.

2 Simple theoretical discussion

2.1 Basic set-up

In the theoretical section I elaborate on the underlying channels through which family policies affect aggregate working hours in a simple macroeconomic model with a representative agent. I recognize that public policies treat groups of households differently, therefore their contribution to changes in working hours are expected to differ as well. The model built here is not a general model and the results are specific to the assumptions made here, but the exercise is interesting and empirically relevant.

I introduce a simple model with two types of households and with a government policy that supports families. Consider a static economy with two types of households: parents (*P*) and nonparents (*N*). They are both assumed to have a log-utility function over the consumption of a market good (c^m), a household good (c^h) and leisure (l). The market good refers to a composite good bought in the market and the household good represents the amount of childcare. Utility functions and budget constraints are as follows.

Parents maximise:

$$\alpha \ln c_P^m + \beta \ln c_P^h + (1-\alpha-\beta) \ln l_P \tag{1}$$

subject to

$$c_P^m = (1-\tau) h_P + T + s \tag{2}$$

$$c_P^h = H_P + g \tag{3}$$

$$l_P = 1 - h_P - H_P \tag{4}$$

Nonparents maximise:

$$\alpha \ln c_N^m + \gamma \ln c_N^h + (1-\alpha-\gamma) \ln l_N \tag{5}$$

subject to

$$c_N^m = (1-\tau) h_N + T \tag{6}$$

$$c_N^h = H_N \tag{7}$$

$$l_N = 1 - h_N - H_N \tag{8}$$

Where

$$\beta > \gamma. \tag{9}$$

τ represents tax wedge[3], T is a lump sum subsidy, s represents cash transfers from the government to parents and g is public provision of childcare. h represents time worked in the market, and H is time spent in childcare at home. α and β are utility parameters on consumption and childcare, respectively. All parameters and variables are assumed to be nonnegative and also $0 < \alpha < 1$, $0 < \beta < 1$ and $0 \leq \gamma < 1$.

As seen from expressions (1) - (9) parents and nonparents differ in certain important aspects. One such difference is in the public policy parameters. Parents receive a cash-transfer s from the government (equations (2) and (6)) capturing child allowances and other direct transfers to parents that exist in many countries. Time constraints (3) and (7) on c^h assume that parents also receive a childcare subsidy g from the government. This is assumed to be a perfect substitute for home provision of childcare and is provided to the households directly, in kind. There is no doubt that this is a simplistic way of entering government support of childcare programmes into the model; it assumes that government simply gives each parental household a certain "amount" g of childcare. In this model, households have no cost of childcare other than the opportunity cost of time, i.e. they don't need to pay for childcare provided in the market (external day care). For alternative specifications, look for example at Ragan (2005) and Rogerson (2007). In their models, childcare can be produced either at home or in the market, with these two inputs being strong substitutes for each other. When the government subsidizes childcare provided in the market, its relative price goes down, and as a consequence individuals spend less time in childcare at home and more time at work and in leisure. In my specification, on the other hand, I assume that there is no market for childcare. Yet the main mechanism is very similar; with the government subsidy g, the government takes care of the children instead of the parents, hence they are able to work more and enjoy more leisure. While my model is simpler with regards to the specification of the government policies, it is more involved in that it introduces two types of households: parents and

[3] Following Nickell (2004) the tax wedge τ incorporates three types of taxes: payroll tax t_1, direct taxes on income t_2 and indirect taxes on consumption t_3. This follows from defining the budget constraint for the market good as $(1+t_3)c^m = (1-t_2)(1-t_1)h$, hence $\tau = 1 - \frac{(1-t_1)(1-t_2)}{(1+t_3)}$. Although I focus only on the labour supply I incorporate the payroll tax t_1 in order to capture potentially important effects of this tax on the demand for labour.

nonparents. In aggregate, the main results of the model presented here and of the models in Rogerson (2007) and Ragan (2005) go in the same direction[4].

Another important difference between nonparents and parents is in their preferences. Condition (9) states that parents derive higher utility from consuming c^h, which is based on a plausible assumption that parents need to spend more time to care for the children. On the other hand, I assume that the utility coefficient α on c^m is the same for both types of households. This is done for two reasons. First, complicating the model by allowing different utility coefficients on market consumption would not add much to answering the questions of interest in this chapter. Second, it is not clear empirically whether parents work more or less than nonparents; in some countries they work more while in other countries they work less[5]. From the assumptions made here it therefore follows that differences in *working time* between parents and nonparents exist solely due to differences in the treatment of the two groups at the hands of public policy rather than due to differences in preferences.

Households maximise utility with respect to hours of work h and hours in childcare at home H, subject to budget and time constraints. See Appendix for first order conditions. Assuming an interior solution, the supply of labour is as follows,

$$h_P = \alpha(1+g) - \frac{(1-\alpha)(T+s)}{(1-\tau)}, \tag{10}$$

$$h_N = \alpha - \frac{(1-\alpha)T}{(1-\tau)}. \tag{11}$$

Assume that everybody works: $h > 0$; this is sensible if we imagine parents and nonparents as households, rather than individuals.

Now I turn to the public sector. The government is assumed to have a balanced budget. Denote the percentage of parental households as δ. The government budget constraint can be written as:

$$\tau(\delta h_P + (1-\delta)h_N) = T + \delta(g+s). \tag{12}$$

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[4] I outline an alternative set up with the market for childcare at the end of the Appendix.

[5] Data on this will be presented later in the paper; see Table 10. Couples normally decide to have children in their prime working age and for this reason they tend to work more than other groups. But even if one compares parents and nonparents of the same age, while it is true that women with young children on average work less, most men with children in fact work more.

The left hand side of (12) represents government tax revenues and the right hand side is government expenditure on a lump sum subsidy and family policies. The lump sum subsidy T will be treated as endogenous, and by combining (10), (11) and (12) can be written as:

$$T=\frac{\alpha(1-\tau)\tau}{(1-\alpha\tau)}-g\delta(1-\tau)-s\delta. \tag{13}$$

From the expression (13) above it is clear that the more government spends on family support policies, given the tax rate, the less remains for the lump sum subsidy. If the government spends a larger share of its expenditures on family services it not only supports young parents to go to work, but further increases the incentives for work, indirectly, by spending less on the lump sum subsidy.

2.2 Effects of family policy on working hours

By combining equation (13) with equations (10) and (11) we can get the labour supplies in reduced form.

Labour supply of parents:

$$h_P=\frac{\alpha(1-\tau)}{(1-\alpha\tau)}+g(\alpha+\delta(1-\alpha))-s\frac{(1-\alpha)(1-\delta)}{(1-\tau)}. \tag{14}$$

Labour supply of nonparents:

$$h_N=\frac{\alpha(1-\tau)}{(1-\alpha\tau)}+g\delta(1-\alpha)+s\frac{(1-\alpha)\delta}{(1-\tau)}. \tag{15}$$

In this model only the supply side of the labour market matters and demand for labour is assumed to be perfectly elastic. Therefore, expressions (14) and (15) are the building blocks of the discussion about how hours of work (labour supply) react to changes in policy parameters. It is also useful to look at the expression for aggregate working hours:

$$h=\delta h_P+(1-\delta)h_N=\frac{\alpha(1-\tau)}{(1-\alpha\tau)}+g\delta. \tag{16}$$

From (14) - (16) it is clear what determines the working hours of households in this model: preferences, taxes, childcare subsidies, cash transfers and the share of people with children in population. Assume that preferences and the population structure remain constant[6]. It can be

[6] In this paper I will for most part assume that the share of parents in population is exogenous. However, it is easy to argue that it is not exogenous. As suggested for example by Ermisch (1989) and McDonald (2000), one

easily shown that the aggregate labour supply (16) and the labour supply of parents (14) decrease with the tax rate. The same goes for the labour supply of nonparents (15) under plausible conditions[7]. The underlying argument behind this result is that higher taxes reduce the marginal benefit from working so that households reduce their hours supplied to the market.

Labour supply is increasing in the amount of the childcare subsidy g in all three expressions (14) - (16), keeping the tax rate constant. It is also straightforward to see from (14) and (15) that the effect of the childcare subsidy is stronger for parents. The labour supply of parents increases because a higher childcare subsidy g results in more spare time that can be supplied to the market. Another channel through which the childcare subsidy g affects working hours is through its effect on the lump sum subsidy T. This is also the mechanism through which the childcare subsidy affects the labour supply of nonparents (15). The government budget constraint (13) dictates that the more money the government spends on childcare subsidies, the less of it is spent on the lump sum subsidy T, which in turn reduces the non-earned income of households and encourages them to work.

Cash benefits to families s have a negative effect on the labour supply of parents (14), but a positive effect on the labour supply of nonparents (15). The reason for this discrepancy is that, while cash benefits *in*crease the non-earned income of parents, they *de*crease the non-earned income of nonparents by lowering the lump sum subsidy T. The latter effect influences both types of households but the direct effect of cash benefits dominates for parents. At the aggregate level (equation (16)) cash benefits don't have any effect on the labour supply; the direct effects on the labour supply of parents are exactly offset by the indirect effects on the labour supply of both groups through the effect on the lump sum subsidy. Effectively, family cash benefit is just a redistribution of a subsidy from one group to another, which in aggregate has no effect on working hours.

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can easily argue that family support increases fertility, and indeed in some countries these measures may have been introduced with this exact objective in mind. Cohen et al. (2007) for example report evidence that the introduction of child subsidies in Israel increased the fertility of married women with two or more children. In this paper I will not take a stand on the direction of causality, hence I will simply report associations between a crude measure of fertility and other variables of interest. I will also implicitly assume that the share of parents in population is constant over time and across countries. It is indeed slow-moving over time, so the first part does not necessarily generate a huge problem, but it is definitely not constant across countries. However, due to endogeneity problems I will leave it out of my empirical analysis. The lack of an appropriate instrument leads me to leave this issue for future research.

[7] Appendix shows this analytically.

Another interesting way in which we can look at these results is to compare the working hours of parents and nonparents. The expression for the difference is:

$$h_P - h_N = \alpha g - s\frac{(1-\alpha)}{(1-\tau)}. \tag{17}$$

From equation (17) it can be nicely seen that increasing the childcare subsidy while holding the tax rate constant increases the working hours of parents relative to the working hours of nonparents, and vice versa for the family cash benefits. The tax wedge reduces the work of parents relative to nonparents. Although this is a simple model, it can be intuitively generalised to real economies. In countries with different family policies, this should be reflected when comparing the hours worked for parents and nonparents. Suppose we add a stochastic component to the above equation (17). In practice, if childcare subsidies g vary across countries more than cash benefits s and more than taxes τ, and if in addition s and g are not too strongly correlated, then g dominates expression (17). This implies that in the data one would expect to see a positive relationship between childcare subsidies and the difference in hours worked between parents and nonparents. In countries with high taxes and high childcare subsidies, nonparents are discouraged from working due to high taxes, while parents are encouraged to work more due to generous childcare subsidies. Hence, according to the model, in such countries, parents should work a lot when compared to nonparents.

Let me summarize this section. According to the model outlined in this chapter, we would expect the effect of childcare subsidies on aggregate working hours 1) to be positive in general, 2) to be positive for both parents and nonparents, but 3) stronger for parents. There should be 4) no aggregate effects of family cash benefits on working hours, 5) a negative effect for parents, and 6) a positive effect for nonparents.

2.3 Effects of family policy on childcare at home

Here I consider the effects of family policies on another use of time – childcare at home. I assume that households *without* children also spend some time in childcare at home, depending on the parameter γ. In this way I capture the fact that nonparents also look after children; i.e. they spend some time looking after grandchildren, nephews and nieces[8]. In fact,

[8] In the year 2000 in Europe, for example, individuals with children of age 0-15 in the household on average spent 21.6 hours per week looking after children, while individuals with no children of age 0-15 in the household spent 1.4 hours per week. The calculation is based on the European Household Panel data for individuals of 16 years of age or more.

the time in childcare at home of parents and nonparents could also be correlated. If parents go to work while members of a nonparent household (i.e. grandparents) look after the children, time spent by parents at home in childcare is substituted for time spent by nonparents in childcare. However, this kind of substitution is not introduced in my model, in the same way that it is not a part of the representative agent model: the increase in childcare subsidies relieves both types of households from childcare. But the effect on nonparents in this model comes solely through the effect of the lump sum subsidy T rather than through any direct relationship between the time spent in childcare of parents and nonparents.

As before we can find reduced form childcare "supply" for parents, nonparents and in aggregate.

Time spent in childcare at home for parents:

$$H_P = \frac{\beta}{(1-\alpha\tau)} - g(1-\beta(1-\delta)) + s\frac{\beta(1-\delta)}{(1-\tau)}. \tag{18}$$

Time spent in childcare at home for nonparents:

$$H_N = \frac{\gamma}{(1-\alpha\tau)} - g\gamma\delta - s\frac{\gamma\delta}{(1-\tau)}. \tag{19}$$

Aggregate time in childcare at home:

$$H = \frac{\alpha\beta + \gamma(1-\delta)}{(1-\alpha\tau)} - g\delta(1-(\beta-\gamma)(1-\delta)) + s\frac{(\beta-\gamma)\delta(1-\delta)}{(1-\tau)}. \tag{20}$$

The results for the time spent in childcare at home basically go in the opposite direction from the results for labour supply. With the rise in the tax rate, each of the three uses of time considered here increases[9], as the tax rate reduces the opportunity cost of non-work. The childcare subsidy g reduces childcare in all the three expressions (18) - (20). The cash benefits to families s increase the childcare for parents (18) and in the aggregate (20), because by increasing the non-earned income of parents they discourage them from working in the market. The cash benefits s, on the other hand, have a negative impact on the childcare of nonparents (19), via their effect on the lump sum subsidy T. These effects are naturally stronger for parents than for nonparents.

To summarize, there is 1) a negative effect of childcare subsidies on aggregate time spent in childcare at home, 2) the effect is negative for both parents and nonparents, but 3) stronger for

[9] The childcare of nonparents decreases with taxes under plausible conditions. See Appendix.

parents. There should be 4) a positive aggregate effect of family cash benefits on childcare time, 5) the effect should be positive for parents, and 6) negative for nonparents.

3 Working hours across countries

I am interested in explaining the variation of working hours across countries, so it is natural to start the empirical investigation by examining these differences at an aggregate level. First I turn to the OECD data on total working hours for the population aged 15-64. My main analysis is not done with the OECD data, but this is the most common source used to report cross-country differences in working hours, which is why I report it here. In Table 1 I show hours worked per person per week in the year 2000, which I further illustrate in Figure 1. Countries with the highest hours of work are the US (27.1), Canada (24.4) and the UK (24.1). Scandinavian countries are mostly in the top half, Sweden (23.6), Denmark (23.0) and Finland (22.7); only Norway (21.0) is in the lower half. Countries with the lowest hours worked are Belgium (18.2), Germany (18.7) and France (18.8). Differences across countries are quite large; the difference between the US at the top and Belgium at the bottom is nearly 9 hours per week.

Statistically, differences in working hours can be decomposed into differences in the employment rates and differences in the length of the working week for employed persons[10]. The length of the working week takes into account the average hours worked in a week by an employed person as well as vacations, public holidays and absences from work (columns (2) and (3) of the Table 1). The employment rates (column (3)) vary from 0.55 in Italy to 0.79 in Norway, a very large difference. Countries such as Italy (0.55), Greece (0.57) and Spain (0.58) have very low employment rates. This is in large part due to the low participation of women in the labour force. The Scandinavian countries, Norway (0.79), Denmark (0.77) and Sweden (0.76), the US (0.76) and UK (0.73), on the other hand, have high employment rates. There are also big differences in hours worked per week by employed persons (column (2)). Interestingly, hours worked per week for employed workers are highest in Greece (40.0), Italy (35.7), the US (35.4) and Spain (34.9). For the three southern European countries this is a consequence of the fact that due to the low participation of women and older workers, employed persons are mostly prime-aged males. Hence the statistic on hours worked is based on a sample where most of the subjects work full-time, whereas in other countries this statistic is reduced by including many part-time employed women and older workers. Countries with

[10] For a more thorough analysis, see Faggio & Nickell (2007).

lowest hours worked by employed persons are the Netherlands (26.3), Norway (26.6) and Germany (28.2). The difference between the top and the bottom is large: almost 14 hours per week. In the remainder of the chapter I will mainly focus on the overall hours worked per person, although in some instances I will also distinguish between the employment decision and hours of work decision.

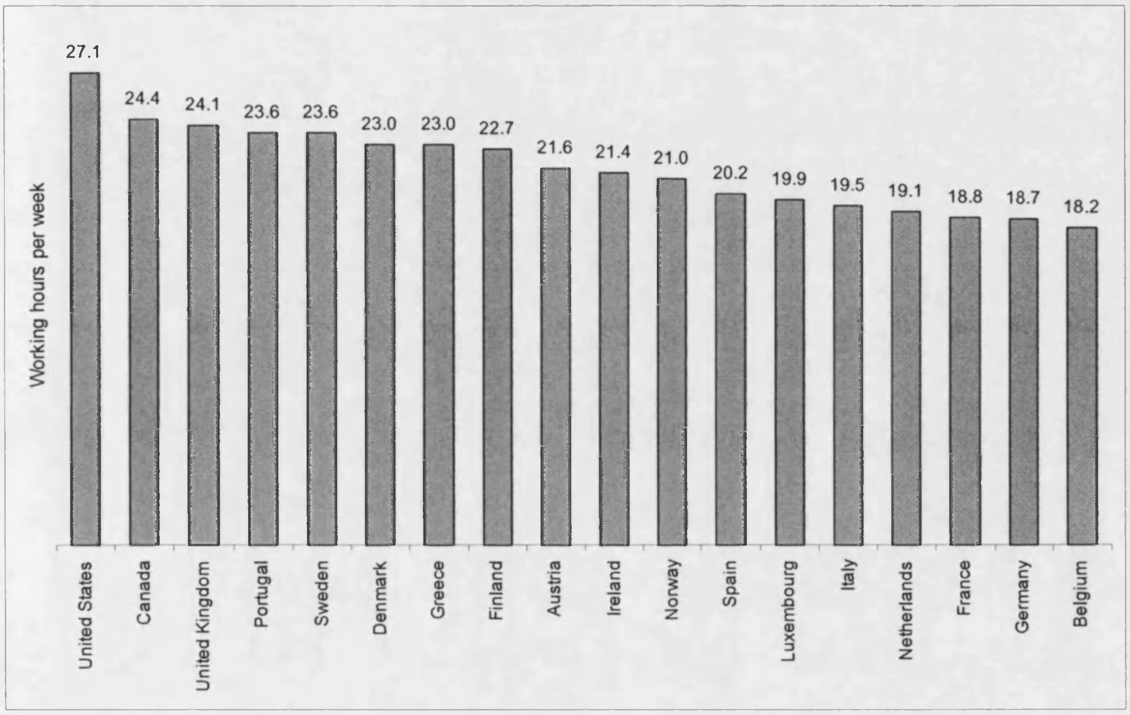
Table 1: Working hours per week per person 15-64 (year 2000)

	(1) = (2)*(3)	(2)	(3)
Country	**Hours per week per person**	**Hours per week (employed)**	**Employment rate**
Austria	21.58	31.38	0.69
Belgium	18.22	29.71	0.61
Canada	24.43	33.97	0.72
Denmark	23.01	29.88	0.77
Finland	22.68	33.64	0.67
France	18.82	30.62	0.61
Germany	18.70	28.23	0.66
Greece	22.96	40.00	0.57
Ireland	21.36	32.46	0.66
Italy	19.52	35.67	0.55
Luxembourg	19.87	31.52	0.63
Netherlands	19.11	26.31	0.73
Norway	21.01	26.54	0.79
Portugal	23.64	32.52	0.73
Spain	20.18	34.91	0.58
Sweden	23.62	31.24	0.76
United Kingdom	24.11	32.85	0.73
United States	27.06	35.40	0.76

Data sources: OECD data on total working hours. Data are for the population 15-64 years of age. Working hours per week per person (1) is the product of the employment rate (3) and the weekly hours worked by employed persons (2). Weekly hours worked by the employed take into account the length of the working week as well as holidays and other days off work.

I now turn to the micro level data obtained from the European Community Household Panel (ECHP). I combine this with the US March Current Population Survey (CPS) from the IPUMS CPS (King et al. (2004)). Data is available at the individual level for people older than 15 years of age at time of survey. To measure working hours I use the variable pe005: Total number of hours worked per week (in main + additional jobs). This variable is based on the following two survey questions: "How many hours per week do you work in your main job, including paid overtime if any?" and "About how many hours per week did you work in your additional job or business? Please give an average figure for the last 4 working weeks." As this variable is only reported for employed individuals, for the rest of the population I assume that they work 0 hours. From the US CPS data I use the variable UHRSWORK: Usual hours worked per week (last year). This variable reports the number of hours per week that respondents usually worked if they worked during the previous calendar year. Individuals were asked this question if they reported working at a job or business at any time during the previous year or if they acknowledged doing "any temporary, part-time, or seasonal work even for a few days" during the previous year. These two combined variables capture the "usual" amount of hours worked per week for an individual. In this sense they capture whether someone participates in the labour force and how much this person usually works. They do not, however, incorporate the number of holidays, vacations and days off work[11].

Figure 1: Working hours per week per person for individuals aged 15-64 (year 2000)



Data sources: OECD data on total working hours. For more details see Table 1.

In Figure 2 I show average working hours per week for individuals aged 16-64 from this data. I choose the age restriction 16-64 in order to be comparable with the OECD data on total working hours reported in Table 1. Working hours on average appear higher than in the OECD data, the reason being that days off work are not counted. The ranking of countries differs from Table 1, but the correlation between the ranks from the OECD data and the micro data is 0.69. The countries with the biggest differences in ranks from one data set to another

[11] It is important to note that respondents report their usual hours of work even if they are temporarily absent from work due to vacation, sickness or injury, maternity leave, bad weather, a lay-off or a labour dispute. Hence, mothers who are on maternity leave report positive hours of work. This is a caveat of the data which unfortunately cannot be overcome.

are Greece (7 places lower), Belgium, Germany (both 6 places lower) and Spain (5 places higher)[12].

4 Family support policies across countries

4.1 Cash benefits and childcare subsidies

I this section I present data on family support policies. OECD Social Expenditure Data (SOCX) provides data on public expenditure on various programmes: old age, survivors, incapacity-related benefits, health, family, active labour market programmes, unemployment, housing and other social policy areas. In what follows I focus on public expenditure under the family programme. More details about the OECD SOCX can be found in OECD (2007) and in the OECD Family Database.

In Table 2 I show family expenditures as a percent of GDP across countries for the year 2000. Public spending on family benefits can be divided into three groups. First, *child related cash transfers* to families with children, column (1). These include items such as child allowances, public income support during parental leave, income support for sole parent families and public childcare support through payments to parents. The second group in column (2) is *public spending on services for families* with children. These comprise of direct financing and subsidising of providers of childcare and early education facilities, public assistance for young people and residential facilities and public spending on family services (centre-based services and home help for families in need). The third group shown in column (3) is *public support for families provided through the tax system*, such as tax exemptions, child tax allowances and child tax credits.

For the purposes of the analysis and to be consistent with the model presented in Section 2 I group these expenditures into two groups: *cash benefits* and *childcare subsidies*, shown on the right hand side of Table 2. Cash benefits capture those transfers from the government that directly increase the non-earned income of households with children. To obtain a measure of cash benefits I combine cash transfers (1) with tax support to families (3). Because tax breaks result in the reduction of the amount households have to pay in taxes to the government, they can be understood as a cash transfer to families. This is an important point because looking

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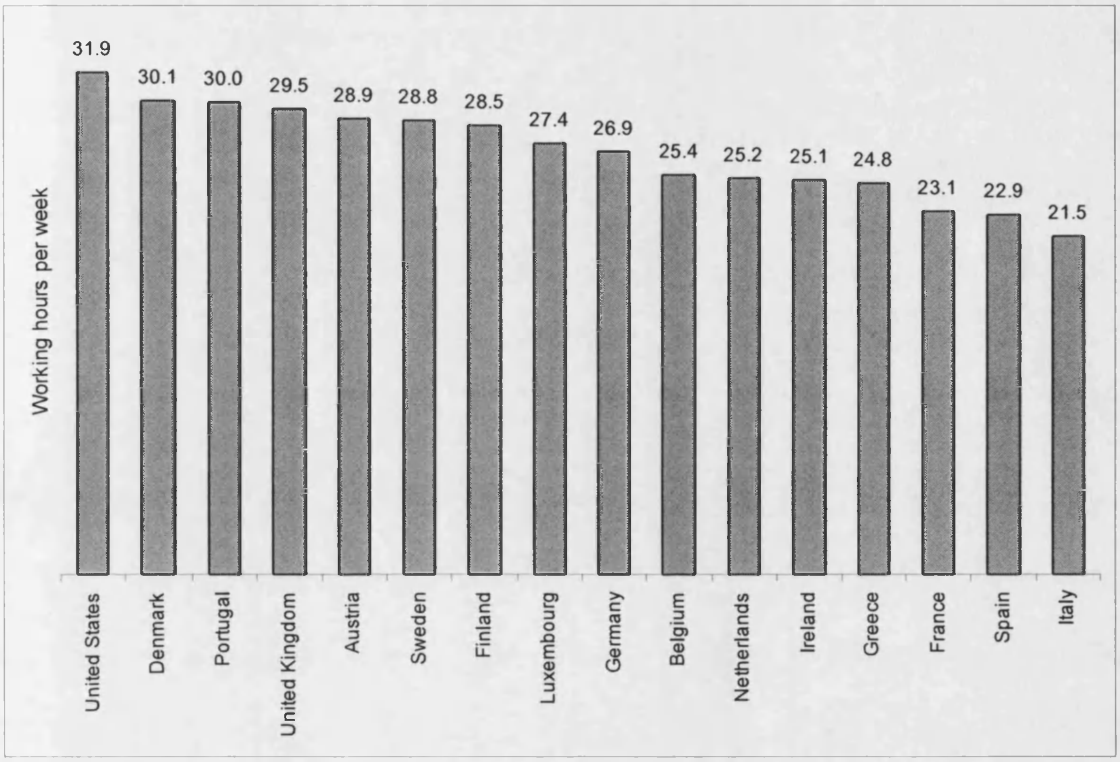
[12] I don't explore the reasons behind such differences, but I conjecture that they cannot be fully explained by the differences in holidays, vacations and days off work. There must be some important differences in data methodology.

only at the cash transfers in column (1) could overstate the actual variation across countries in family cash benefits. In the US, for example, direct cash transfers to families are very low, but the tax system provides certain benefits to families. On the other hand, in Scandinavian countries there is an extensive support to families via cash transfers, but tax reliefs are relatively low.

Figure 2: Working hours per week per person for individuals aged 16-64 from the micro data (year 2000)



Data Sources: European data is from the European Community Household Panel (ECHP). As a measure of working hours I use the variable pe005: Total number of hours worked per week (in main + additional jobs). This variable is based on the following two survey questions: "How many hours per week do you work in your main job, including paid overtime if any?" and "About how many hours per week did you work in your additional job or business? Please give an average figure for the last 4 working weeks." It is only reported for individuals active in the labour market, hence for the rest I assume that they work 0 hours. The US data is from the US March Current Population Survey (CPS) from IPUMS CPS (King et al. (2004)). For the US I use the variable UHRSWORK: Usual hours worked per week (last year). It reports the number of hours per week that respondents usually worked if they worked during the previous calendar year. Individuals were asked this question if they reported working at a job or business at any time during the previous year or if they acknowledged doing "any temporary, part-time, or seasonal work even for a few days" during the previous year. The data is for the year 2000 (wave 7 of the ECHP).

Table 2: Public family spending, % GDP (year 2000)

	(1)	(2)	(3)	= (1) + (3)	≈* (2)
country	**Cash**	**Services**	**Tax breaks**	**Cash benefits**	**Childcare subsidies**
Austria	2.37	0.56	0.02	2.39	0.34
Belgium	1.76	0.83	0.55	2.31	0.59
Canada	0.77	0.20	0.08	0.85	0.20
Denmark	1.49	2.19	0.00	1.49	2.05
Finland	1.70	1.35	0.00	1.70	1.19
France	1.47	1.54	0.77	2.24	1.19
Germany	1.16	0.78	1.01	2.16	0.40
Greece	0.74	0.39	0.00	0.74	0.15
Ireland	1.64	0.19	0.07	1.71	0.00
Italy	0.54	0.61	0.00	0.54	0.55
Luxembourg	2.74	0.48	0.00	2.74	0.47
Netherlands	0.72	0.73	0.44	1.17	0.73
Norway	1.87	1.93	0.12	1.99	1.43
Portugal	0.65	0.43	0.17	0.83	0.31
Spain	0.30	0.56	0.06	0.36	0.40
Sweden	1.52	1.74	0.00	1.52	1.30
United Kingdom	1.87	0.71	0.31	2.18	0.56
United States	0.11	0.63	0.83	0.94	0.38
Average	**1.30**	**0.88**	**0.25**	**1.55**	**0.68**
Coefficient of variation	**0.56**	**0.67**	**1.32**	**0.47**	**0.79**

Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1). Columns (1)-(3) are as reported in the data sources; columns on the right hand side are as shown. *Cash transfers* to families (1) include items such as child allowances, public income support during parental leave, income support for sole parent families and public childcare support through payments to parents. *Services* (2) comprise of direct financing and subsidising of providers of childcare and early education facilities, public assistance for young people and residential facilities, public spending on family services (centre based services and home help for families in need). *Tax breaks* (3) represent tax exemptions, child tax allowances and child tax credits.
Childcare subsidies exclude those services that do not support childcare; such items include for example assistance for pupils and youths, youth centres, family accommodation benefits, transport subsidies, holidays for schoolchildren, school meals etc.

However, it is not always clear that tax breaks fall into the category of cash benefits. In some countries tax deductions can only be claimed based on the family's expenses for external childcare. In such cases these should perhaps be treated as childcare benefits. With available data it is not possible to make such a distinction. Nevertheless, as argued in Immervoll & Barber (2005), page 17, support through the tax system blurs the perceived link between childcare expenses and support payments. Whereas parents need to pay for childcare today, they get reimbursed through the tax system only after the end of the tax year. Therefore the perceived link between expenses and benefits is weakened and tax benefit when received by the family may be seen as a windfall rather than a consequence of their childcare choices. The data on tax breaks for families is only available for the year 2003, so I assume that the ratio of tax breaks over cash transfers remains constant over time in a given country[13].

Childcare subsidies capture transfers from the government that provide services to parents and effectively act as a substitute for parental childcare. Childcare subsidies include public spending on pre-primary education, data on which is available only from the year 1998 for most countries. Data on childcare subsidies before the year 1998 are thus not comparable across countries. Another issue with the reported measure of public support for services to families (column (2) of Table 2) is that it includes items such as assistance for pupils and youths, youth centres, family accommodation benefits, transport subsidies, holidays for schoolchildren and school meals. These do not represent a substitute for the childcare of parents. As much as the information in the OECD SOCX dataset allows, I exclude these items from the measure of childcare benefits. I consider the new measure of childcare subsidies as a more appropriate one and use it in the analysis that follows. My results are similar regardless of which of the two measures I choose. Note also that the correlation coefficient between the two measures is 0.97.

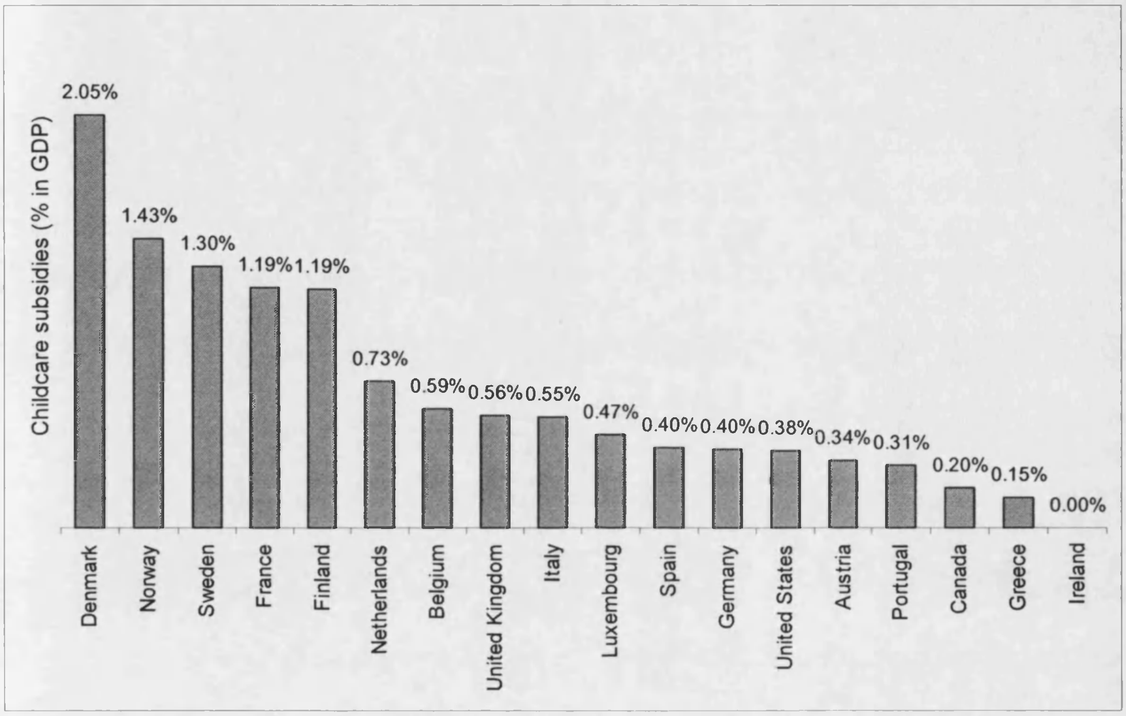
Let me now turn to the information contained in Table 2. From the last two columns we can see that the variation in childcare subsidies is considerably higher than the variation in cash benefits. Figure 3 and Figure 4 further illustrate the variation in public child support across countries. Childcare subsidies (Figure 3) tend to be especially high in the Scandinavian countries: Denmark (2.05%), Norway (1.43%), Sweden (1.30%) and Finland (1.19%)[14]. The share of childcare subsidies in GDP is also very high in France (1.19%). Countries with the lowest share of childcare subsidies in the GDP are Ireland (0.00%), Greece (0.15%) and Canada (0.20%). In Figure 4 I show family cash benefits, which are the highest in Luxembourg (2.74%), Austria (2.39%) and Belgium (2.31%), and the lowest in Spain (0.36%), Italy (0.54%) and Greece (0.74%).

I now turn my attention to the relationship of the two family support policies with each other. Figure 5 plots a scatter diagram of cash benefits and childcare subsidies. There is a weak positive correlation (0.21), not significantly different from zero. Countries can be divided roughly in three groups according to the combination of childcare subsidies and cash benefits. The first group consists of countries with high childcare subsidies and a medium to high level of cash benefits. This group comprises of the Scandinavian countries (Denmark, Norway,

[13] Time variation in these measures is in general low hence the error thus committed should not be too large.
[14] This empirical observation was the basis of discussion in the papers by Ragan (2005) and Rogerson (2007).

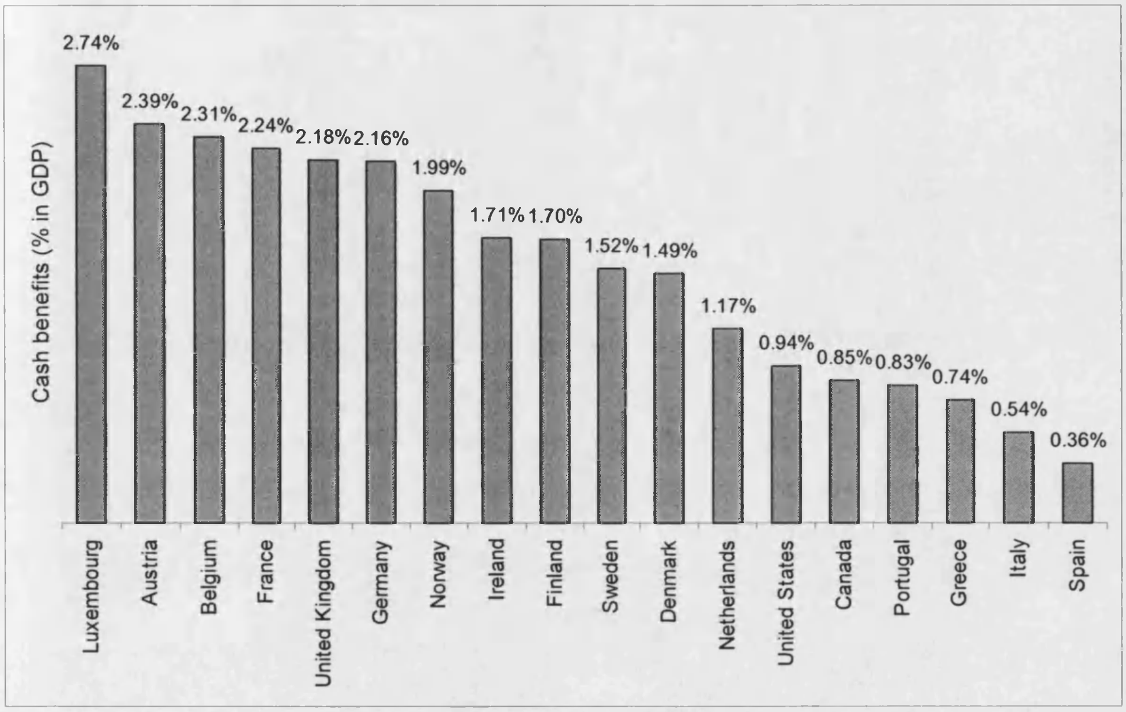
Sweden, and Finland) and France. The second group consists of countries with low childcare subsidies but relatively high cash benefits: Luxembourg, Belgium, Austria, UK, Germany and Ireland. And lastly, the third group consists of countries with low subsidies and low cash benefits: Spain, Italy, Greece, US, Portugal, Canada and the Netherlands.

Figure 3: Childcare subsidies, % in GDP (year 2000)



Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1).

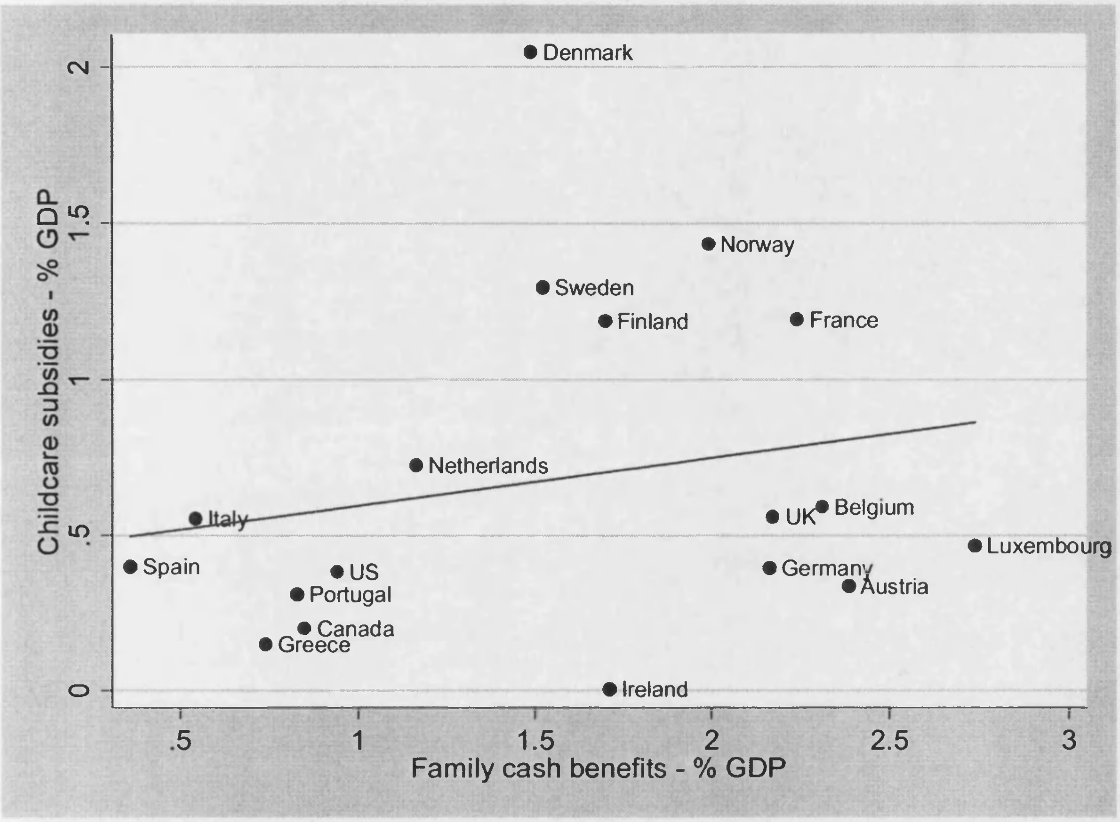
Figure 4: Family cash benefits, % in GDP (year 2000)



Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1).

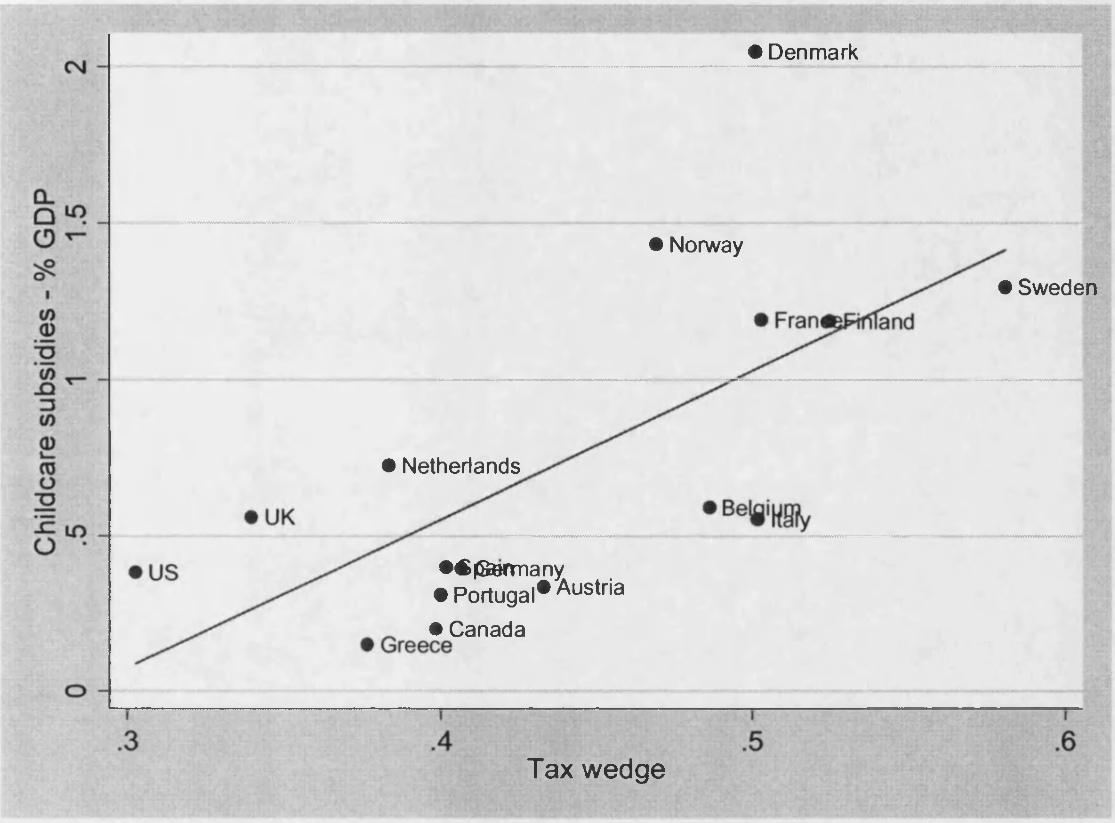
Another important policy variable in the model presented above is the tax rate. In Figure 6 and Figure 7 I therefore show scatter plots of family support policies against the tax wedge. The average tax wedge is calculated as explained in the footnote 3, based on the CEP-OECD data on labour institutions described in William Nickell (2006). As Greece is not included in the CEP-OECD data, I calculated the tax wedge directly from the OECD data as explained in Nickell (2006). Figure 6 reveals a strong positive relationship between taxes and childcare subsidies across countries; the correlation coefficient is 0.66 and statistically significant. The Scandinavian countries and France stand out as having both high tax rates and high childcare subsidies. The US and UK have low tax rates and medium levels of childcare subsidies. On the other hand, Figure 7 shows that there is no significantly positive correlation between taxes and family cash benefits. In terms of cash benefits, the Scandinavian countries do not stand out. Together these two figures suggest a potential for family policies to dampen the effect of taxes on working hours.

Figure 5: Childcare subsidies and family cash benefits, % in GDP (year 2000)



Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1). There is a positive correlation (0.21), not significantly different from zero.

Figure 6: Childcare subsidies (% in GDP) and tax wedge (year 2000)

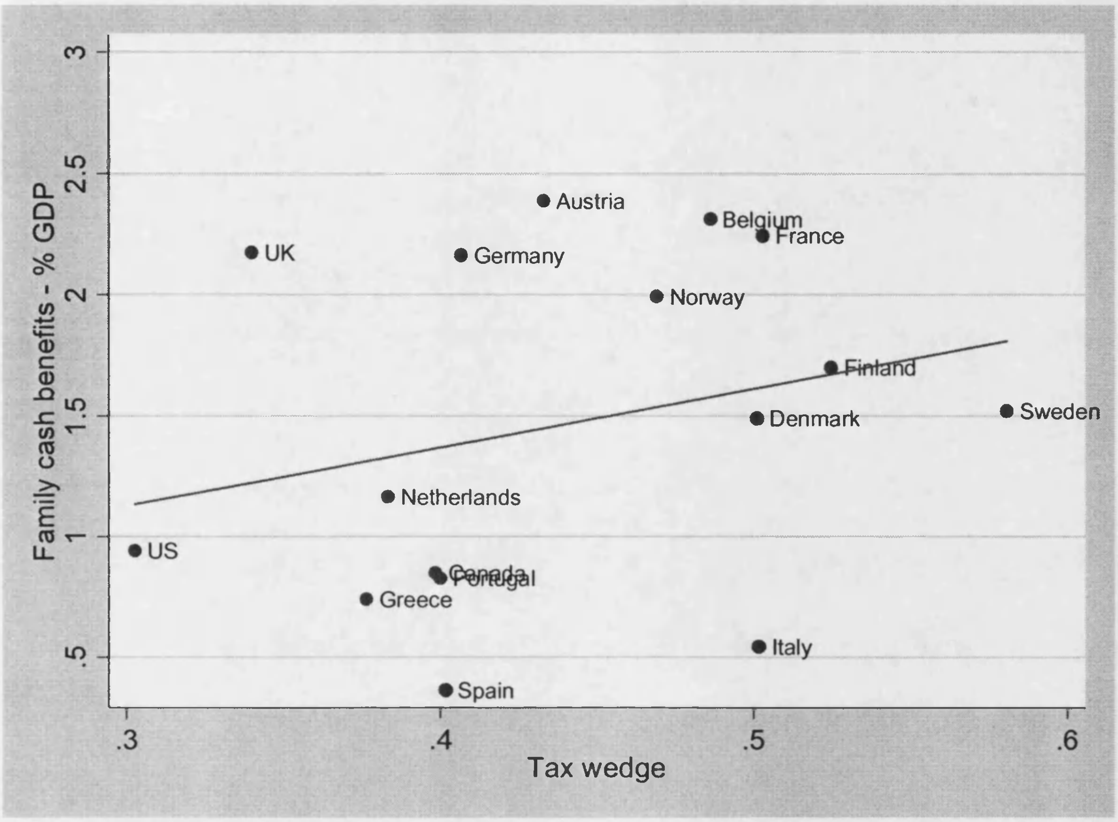


Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1). Taxes are obtained from the CEP-OECD dataset, described in William Nickell (2006). There is a statistically significant positive correlation (0.66).

4.2 Children in formal care

In this section I look at the percentage of children enrolled in external childcare and relate it to family policies. I do this in order to see whether family policies have the effects that one would intuitively expect and also to get a sense of the degree to which the empirical measures used in this chapter are informative. Childcare subsidies, for example, directly affect both the demand for external childcare (by subsidising the fees households pay for external care) and the supply of external childcare (via the public provision of childcare-centres or subsidies to childcare providers). Thus childcare subsidies are expected to have a strong impact on the proportion of children enrolled in some type of external care.

Figure 7: Family cash benefits (% in GDP) and tax wedge (year 2000)



Data sources: OECD Social Expenditure database (SOCX 1980-2003) and OECD Family Database (PF1). Taxes are obtained from the CEP-OECD dataset, described in William Nickell (2006). There is a positive correlation (0.26), not significantly different from zero.

In Table 3 I present data on the use of external childcare, and in Figure 8 I plot the relationship between external childcare and childcare subsidies across countries. The first two columns in Table 3 show the enrolment rates of children under six in childcare and early education services. The data is obtained from the OECD Family Database. Enrolment rates for those aged 0-2 years primarily refer to formal childcare arrangements such as group care in childcare centres, registered childminders based in their own homes and care provided by a carer at home who is not a family member. Enrolment rates for those aged 3-5 years refer to those enrolled in formal pre-school services, and, in some countries, 4 and 5 years old in primary schools. Where children are enrolled in more than one part-time programme, the issue of double counting arises. This leads to overestimated enrolment rates. As a result, reported enrolment in some countries can exceed 100% and the data needs to be interpreted with caution.

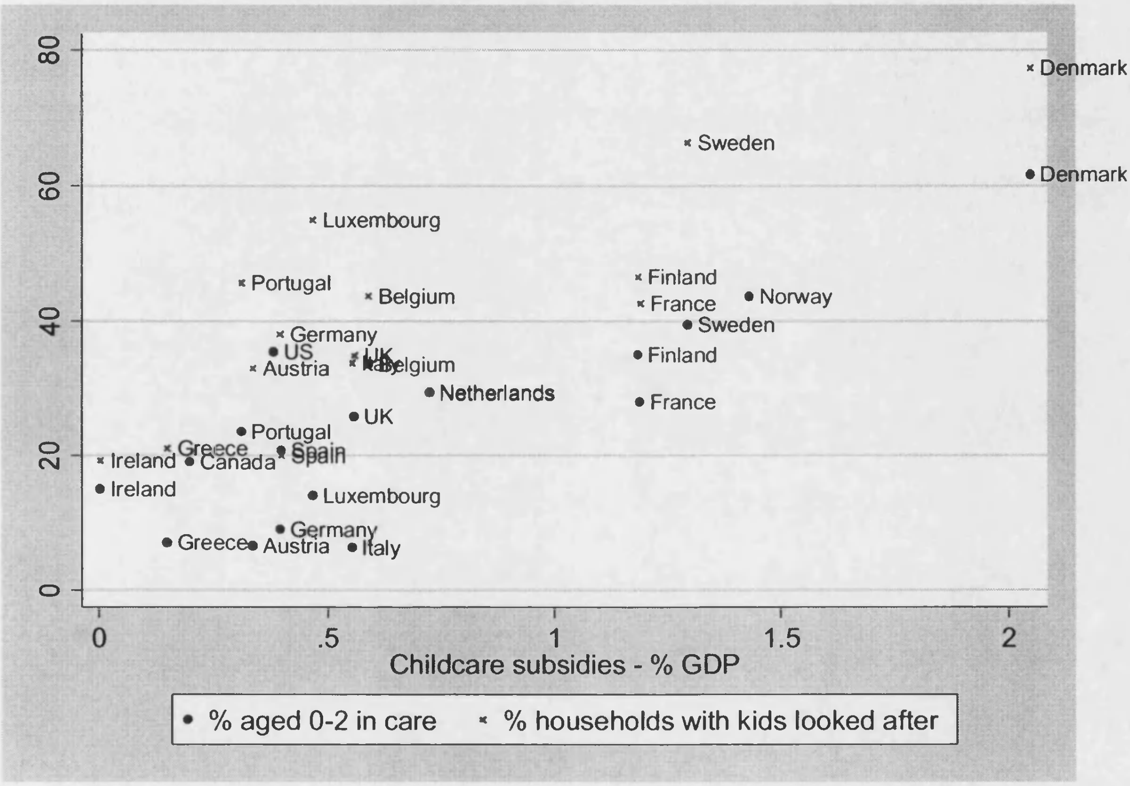
Table 3: Use of formal childcare

	OECD Family Database Enrolment rates of children under six in childcare and early education services (2003/04)		ECHP % of households having their kid looked after on a regular basis (2000)
Country	**0 - 2 years**	**3 - 5 years**	**0-11 years**
Austria	6.6	74.0	33.0
Belgium	33.6	99.6	43.7
Canada	19	na	na
Denmark	61.7	89.7	77.4
Finland	35	46.1	46.6
France	28	101.9	42.7
Germany	9	80.3	38.1
Greece	7	46.8	21.0
Ireland	15	68.2	19.2
Italy	6.3	100.3	33.8
Luxembourg	14	72.3	55.0
Netherlands	29.5	70.2	29.3
Norway	43.7	85.1	na
Portugal	23.5	77.9	45.7
Spain	20.7	98.6	20.0
Sweden	39.5	86.6	66.4
United Kingdom	25.8	80.5	34.9
United States	35.5	62.0	na
Correlation with childcare subsidies	**0.83***	**0.28**	**0.81***

Data sources: OECD Family Database (PF11). Canada and Germany (2001); France (2002); Greece, Luxembourg and Norway (2003); Denmark and the United States (2005). Enrolment rates for those aged 0-2 years primarily refer to formal childcare arrangements such as group care in childcare centres, registered childminders based in their own homes and care provided by a carer at home who is not a family member. Enrolment rates for those aged 3-5 years refer to those enrolled in formal pre-school services, and in some countries 4 and 5 years old in primary schools. Where children are enrolled in more than one part-time programme, the issue of double counting arises. This leads to overestimated enrolment rates (enrolment in some countries > 100%).
ECHP (European Community Household Panel), year 2000. Calculations based on the survey question at the level of household: "Are any of the children (0-11) in this household looked-after on a regular basis by someone other than their parent or guardian, whether at home or outside such as at a crèche or kindergarten?" Based on all households with at least one child aged 0-11.
* denotes significantly different from zero at 5%.

Differences in enrolment rates across countries are large. For 0-2 year-olds, enrolment rates range from 6.3% in Italy to 61.7% in Denmark. Variation in the enrolment of 3-5 year-olds is much lower, mainly due to the fact that most children of this age are in some form of pre-school service. The lowest enrolment is in Finland (46.1%) and the highest in France (101.9%). The data, however, does not account for other important differences in childcare patterns. For example, the number of hours children typically spend in formal care varies across countries. According to Immervoll & Barber (2005), in some cases, having such information would further amplify the observed variation across countries. There are other issues with the comparability of the data across countries (see OECD Family Database (PF11)). Enrolment rates suffer from underreporting in countries where childcare and early education is to a large extent provided by the local government (Canada, Ireland and the US); there are also important differences in parental leave arrangements across countries (these can be considered as family cash benefits), which influence the extent to which children are cared for by external institutions. Countries also differ in the availability of informal care by other family members (i.e. grandparents).

Figure 8: Use of formal childcare and childcare subsidies (% in GDP)



Data sources: OECD Family Database, ECHP (European Community Household Panel) and OECD Social Expenditure database (SOCX 1980-2003). For more details see notes under Table 3.

In the last column of Table 3 I report the % of households with at least one child aged 0-11 that have their children looked after on a regular basis. Data is based on the ECHP survey question: "Are any of the children (0-11) in this household looked after on a regular basis by someone other than their parent or guardian, whether at home or outside such as at a crèche or kindergarten?" Variation is again very large, ranging from 19.2% of households in Ireland to 78.2% of households in Denmark. It is not clear, however, whether the question refers to formal or informal child care; answers most probably take into account both types of childcare.

Figure 8 shows a scatter plot of formal childcare use versus childcare subsidies. I use two different measures of formal childcare use: enrolment of 0-2 year olds based on the OECD data and the % of households with children cared for based on the ECHP. There is a clear positive correlation between childcare subsidies and use of external childcare (see also bottom of Table 3). Note, however that the use of childcare for children of age 3-5 years is not significantly correlated with childcare subsidies. Since in many countries most of the 3-5 year-olds are enrolled in some form of external care, it is possible that a high share of the observed variation for this group is due to measurement error. On the other hand, parents of older children in general participate in the labour force more than parents of young children. Hence, intuitively, what should matter for the explanation of the hours worked across countries with respect to public family policies are the employment and work decisions of parents with very small children.

Table 4: Relationship between the use of formal childcare and family policy variables (year 2000)

Dependent variable:	% aged 0-2 in care	% aged 3-5 in care	% households with kids looked after+
	(1)	(2)	(3)
Childcare subsidies - % GDP	23.647	8.748	23.657
	(4.022)***	(8.251)	(3.268)***
Family cash benefits - % GDP	-2.261	0.430	3.877
	(2.996)	(6.175)	(2.502)
Constant	12.628	71.956	17.139
	(5.288)**	(11.406)***	(5.508)***
Observations	18	17	59
R-squared	0.70	0.08	0.75

Standard errors in parentheses, +pooled OLS: period 1998-2001, standard errors clustered by country, with 15 clusters. * significant at 10%; ** significant at 5%; *** significant at 1%

So far the analysis of the use of external childcare across countries has been descriptive. Therefore, in Table 4 I report results from regressing each of the three measures of the use of formal childcare on the childcare benefits and the cash benefits. Coefficients in the regression with the % of 3-5 years old in formal care (column (2)) are not significant. However, the results from regressions with the % of 0-2 year olds in formal care (column (1)) and the % of households with children looked after (column (3)) show that the childcare subsidies have a strong positive effect on the use of formal childcare. The effects of family cash benefits, on the other hand, are not statistically significant. To sum up the discussion in this section, I conclude that there is evidence that my measure of childcare subsidies is capturing something sensible, and that childcare subsidies have a significant effect on the behaviour of households.

5 Effects of family support policies on working hours

5.1 Model calibrations: What is the magnitude of the effects?

In light of the data on family policies presented above, it would be interesting to check the magnitudes of the effects on working hours implied by the model from Section 2. The country with the highest childcare subsidies, Denmark, devotes 2.05% of its GDP for this measure. Is this high enough to result in a significant effect on aggregate working hours? What about other countries, with smaller share of family policies in their GDP? If the implied effects on working hours are found to be small, then it is hard to argue that family policies can contribute much to explaining differences in working hours across countries.

The results of this exercise are presented in Table 5. I use the OECD data on working hours and the OECD SOCX data on childcare subsidies and cash benefits in order to obtain corresponding values for h, τ, g and s (columns (1) - (4)). To obtain a value of the utility parameter α, I calibrate the model using equation (16) with the US data and obtain a value of $\alpha = 0.35$. I use this value for all countries. Using the US CPS data, I also obtain a value of $\delta = 0.39$, which I also use for all countries. Columns (5) - (9) of Table 5 report the % change in working hours implied by the introduction of family policies, based on equations (14) - (16). Changes are computed by comparing working hours in two settings: in the first setting I set the family policy parameters equal to their actual values, whereas in the second setting I set them equal to zero. I report results for 6 cases choosing two countries from each group based on Figure 5. Sweden and Denmark represent the high subsidy - high cash benefits group, Belgium and UK represent the low subsidy - high cash benefits group, and Italy and US represent the low subsidy - low cash benefits group. Note that in all cases, one of the two countries has relatively high tax rates and the other one relatively low tax rates.

On aggregate (column (5)), family policies do not seem to have a very strong effect on working hours. The biggest effect, as expected, is in Scandinavia, where family policy variables increase working hours for 2.25% in Denmark and 1.68% in Sweden compared to a situation with no family policies, but keeping tax rates constant. In contrast, Ragan (2005) in Tables 1, 5 and 6, evaluates corresponding effects to be equal to 16.7% and 25.0%, respectively. Rogerson (2007) in Table 6, evaluates that the public provision of family services increases working hours by 8.2% for Scandinavian levels of policy parameters. Although their models are more complex, they don't recognise that family policies *directly* influence only a certain proportion of the population[15]. Furthermore, in their measure of public support to family services they include subsidies for elderly care services; this is not justified and overstates the effect of public subsidies on working hours. ECHP data for the year 2000 reveal that in Europe persons of age 16-55 on average spend 0.7 hours per week looking after a person with special needs without pay, while persons older than 55 years spend 1.6 hours per week. Hence this activity does not take much time for an average person and furthermore it is concentrated among older people, who for the most part do not participate in the labour market regardless of public support for families.

Table 5: Calibration of the model: magnitude of the effects of family policies on working hours

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					Aggregate	Parents		Nonparents	
Country	**h**	**τ**	**g**	**s**	**effect of including g and s**	**effect of including g and s**	**effect of including g only**	**effect of including g and s**	**effect of including g only**
Sweden	0.24	0.58	0.0080	0.0094	1.68%	-2.16%	2.61%	4.15%	1.09%
Denmark	0.23	0.50	0.0123	0.0090	2.25%	0.17%	3.49%	3.59%	1.46%
Belgium	0.19	0.49	0.0028	0.0110	0.50%	-3.09%	0.78%	2.81%	0.33%
UK	0.25	0.34	0.0035	0.0137	0.52%	-2.29%	0.81%	2.32%	0.34%
Italy	0.20	0.50	0.0028	0.0028	0.52%	-0.23%	0.80%	1.00%	0.34%
US	0.28	0.30	0.0027	0.0067	0.38%	-0.77%	0.59%	1.13%	0.25%

Time frame for calculations of h is 14 hours a day (as in Ragan (2005) and Prescott (2004)), $h = (\text{OECD working hours})/(14*7)$. Policy parameters are based on data sources described above (OECD, OECD SOCX). In the model h represents the value of the whole economy (= gdp), hence the real value of what government gives for childcare subsidies is $totg = (\text{childcare subsidies as \% GDP})*h$. This only goes to parents, hence $g = totg / \delta$. s is calculated in the same way.

In the calculations I compare values for h in the model where g (and s) is set to their actual values, with h from the model where g and s are set to zero, based on equations (14)-(16). I set $\alpha = 0.35$, calibrated from the US data using equation (16), and $\delta = 0.39$, calculated from the US CPS.

In columns (6) and (7) of Table 5 I show the effects of family policies on parents. According to my model, on average family policies actually reduce the working time of parents due to the negative effect of cash benefits. This becomes clear after inspecting column (7), where cash benefits are set to zero and only the effect of childcare subsidies is considered; the effect on working time is positive, especially so for Scandinavian countries. Columns (8) and (9) report effects for nonparents. The total effect on this group is quite large, as both subsidies and cash benefits work in the same direction (via reducing the lump sum subsidy). However, as expected, when looking at the effects of childcare subsidies only, the effects are smaller than for parents. To summarise, in the aggregate, childcare policies do not seem to have very

[15] The Ragan (2005) and Rogerson (2006) models also don't account for the effects of cash benefits. However, in aggregate cash benefits don't have any effect in my model either, as they just redistribute a cash subsidy from one group to another.

strong effects on working hours. Even for countries with generous public policies such as Scandinavian countries, the effects on working hours are only around 2%. This does not seem large enough to contribute greatly to the explanation of working-hours differences across countries.

5.2 OECD total working hours and family policy

I now turn to the empirical analysis of the question of whether family policies can explain the variation in working hours across countries and whether the effects are in accordance with the theory outlined in Section 2. In Table 6 I report regressions of OECD working hours on the family policy variables and other controls. Due to the low variation of policy variables over the short period of time, I use cross-sectional variation, with 16 countries included in the analysis. However, to increase precision I use a pooled OLS regression over the period 1998-2003 with reported standard errors clustered by country. I report three specifications. In column (1) I report a specification with childcare subsidies, family cash benefits and the tax wedge on the right hand side only. This specification refers to the argument by Prescott (2004) augmented by the public policy arguments of Ragan (2005) and Rogerson (2007). In the second specification I add the aggregate unemployment rate, which accounts for any involuntary unemployment. In countries with a higher unemployment rate, working hours may be lower due to fewer opportunities to find jobs. In the third specification, following the empirical analysis in Alesina et al. (2005), Tables 9-10, and Faggio & Nickell (2007), Table 16, I include additional institutional indicators from the CEP-OECD dataset: union density and degree of employment protection. They control for differences in the rigidity and nature of labour markets across countries. As argued in Alesina et al. (2005), trade unions respond to adverse shocks by trying to protect employees, pressing for work sharing as well as employment protection more generally, which tends to reduce overall working hours. Therefore, one should control for these institutional characteristics. The regression analysis throughout this chapter will be based on these three specifications.

In column (1) of Table 6 the direction of the coefficients is more or less consistent with the theory; childcare subsidies have a significant positive effect, cash benefits have an insignificant (negative) effect and taxes have a negative though insignificant effect. However, after including more controls in columns (2) and (3), the results become by and large insignificant and inconsistent with the theory. After including the aggregate unemployment rate, the positive effect of childcare subsidies becomes insignificant, whereas the negative cash benefits coefficient becomes statistically significant. After including union density and employment protection measures, all three policy variables are insignificant. The effect of childcare subsides is close to zero, the effect of family cash benefits is negative but insignificant, and the effect of the tax wedge is negative, but too imprecisely measured, hence insignificant. Also all three included controls have statistically insignificant coefficients. It appears, therefore, that the OECD data does not have enough power to yield any firm conclusions on family policies and their role in explaining working-hours differences across countries.

Table 6: OECD total working hours and family policy: regression analysis (1998-2003)

Dependent variable: Working hours per week			
	(1)	(2)	(3)
childcare subsidies - % GDP	1.892	1.152	0.199
	(0.983)*	(0.836)	(1.070)
family cash benefits - % GDP	-1.018	-1.414	-1.344
	(0.784)	(0.702)*	(0.779)
tax wedge	-17.363	-9.553	-8.080
	(11.640)	(11.364)	(11.912)
unemployment rate		-0.295	-0.200
		(0.167)*	(0.171)
union density			0.033
			(0.039)
employment protection			-3.544
			(2.411)
Constant	29.079	28.874	29.499
	(4.907)***	(5.116)***	(3.955)***
Observations	82	80	63
No. of clusters/countries	16	15	15
R-squared	0.25	0.34	0.55

Robust standard errors in parentheses (clustered by country), * significant at 10%; ** significant at 5%; *** significant at 1%

5.3 Regressions of working hours on policy variables

I now turn to the micro level data where I will focus on comparisons of two groups: people with children and people with no children. For this I combine data from various sources. Data on hours worked together with personal and household characteristics are obtained from the European Community Household Panel (ECHP) and the US March Current Population Survey (CPS). The data set includes 15 EU countries and the US. The period of analysis will mainly be for the years 1998-2001, where policy variables and household panel data are consistently available. I supplement the individual level data with country level variables measuring family-support policies from the OECD Social Expenditure Data (SOCX) and measures of taxes and labour market institutions from the CEP-OECD Institutions Dataset (William Nickell (2006)).

First let me briefly discuss the regression estimation. I pool together the whole available period 1998-2001 in order to increase efficiency. To be on the conservative side, however, I cluster standard errors by country, despite having multiple observations over time for each country. The main variation exploited in this analysis is thus cross-country variation, with 14 countries[16]. Due to the short time period and the slow movement of policy variables over time, I do not rely on fixed effects regressions. I report weighted regressions, using weights as provided in the original data sets. However, in order for each country to have the same weight, regardless of the sample size, I convert the weights so that the summation of weights equals to 1.0 for each country in each year.

Table 7 shows the results from regressing individual working hours on policy variables for individuals older than 15 years. All regressions control for an individual's age and age squared. Column (1) includes the main three fiscal policy parameters only: the childcare subsidies, the family cash benefits and the tax wedge. The results are consistent with the theory. Childcare subsidies have a significant positive effect and the tax wedge has a significant and negative effect. Family cash benefits have an insignificant effect on working hours. The size of the coefficients, however, decreases after including the aggregate unemployment rate, as reported in column (2). The effects of childcare subsidies still remain statistically significant and positive, but the magnitude of the effect is reduced. In column (3) I report my preferred specification with the unemployment rate and the two measures of the rigidity of labour market institutions. The coefficients for childcare subsidies and cash benefits are both insignificant. The tax wedge has statistically significant negative effect, as expected. The same goes for the aggregate unemployment rate, which raises my confidence in this specification. Union density has a significant positive effect, corroborating the results in Faggio and Nickell (2007) and Bowles and Park (2005), but going against the story in Alesina et al. (2005). The effect of employment protection is not statistically different from zero.

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[16] Only 14 instead of 16 countries are included in the regression analysis because data on tax rates for Ireland and Luxembourg are not available.

Table 7: Working hours and family policy: regressions on the sample of age 16 and more (1998-2001)

Dependent variable: Working hours per week			
	(1)	(2)	(3)
childcare subsidies - % GDP	2.633	1.443	0.413
	(0.972)**	(0.792)*	(0.925)
family cash benefits - % GDP	0.664	-0.078	-0.098
	(0.868)	(0.743)	(0.566)
tax wedge	-24.792	-12.232	-29.249
	(9.996)**	(8.314)	(10.652)**
unemployment rate		-0.450	-0.297
		(0.172)**	(0.150)*
union density			0.084
			(0.026)***
employment protection			0.190
			(2.793)
age	1.779	1.772	1.777
	(0.144)***	(0.151)***	(0.151)***
age^2	-0.022	-0.022	-0.022
	(0.001)***	(0.001)***	(0.001)***
Constant	3.110	3.082	6.884
	(5.426)	(5.136)	(4.807)
Observations	829502	810698	810698
No. of clusters/countries	14	13	13
R-squared	0.26	0.27	0.27

Robust standard errors in parentheses (clustered by country), * significant at 10%; ** significant at 5%; *** significant at 1%

To further explore what is driving the results, I decompose working hours into two components: first, the decision of whether or not to participate in the labour market, and second, the number of hours to work per week conditional on participating in the labour market. In Table 8 I report three different specifications. The "total effects" specification, where everyone is included in the sample, active as well as inactive people, and where working hours are on the left hand side, is reported in columns (1) and (4). Results of this specification are equivalent to the results reported in columns (1) and (3) of Table 7. In columns (2) and (5) of the Table 8 I report regressions with hours worked conditional on working on the left hand side. Hence, only employed people are included. The last specification, in columns (3) and (6), estimates the probability of working, with a categorical variable indicating whether an individual is employed (= 1) or not (= 0) on the left hand side. Marginal effects from probit regressions are reported.

Table 8: Participation, working hours of employed people and family policy: regressions on the sample of age 16 and more (1998-2001)

Dependent variable:	Working hours per week		Employed dummy	Working hours per week		Employed dummy
Sample:	All	employed	all	all	employed	all
	OLS		Probit+	OLS		Probit+
	(1)	(2)	(3)	(4)	(5)	(6)
childcare subsidies - % GDP	2.633	-2.711	0.159	0.413	-2.412	0.130
	(0.972)**	(0.889)***	(0.029)***	(0.925)	(1.036)**	(0.030)***
family cash benefits - % GDP	0.664	-0.462	0.049	-0.098	-0.063	0.025
	(0.868)	(0.552)	(0.024)**	(0.566)	(0.576)	(0.016)
tax wedge	-24.792	12.735	-1.085	-29.249	-4.074	-0.526
	(9.996)**	(7.597)	(0.256)***	(10.652)**	(10.387)	(0.327)
unemployment rate				-0.297	0.240	-0.014
				(0.150)*	(0.174)	(0.003)***
union density				0.084	0.045	-0.001
				(0.026)***	(0.029)	(0.001)
employment protection				0.190	0.986	-0.010
				(2.793)	(2.646)	(0.070)
age	1.779	1.042	0.061	1.777	1.046	0.060
	(0.144)***	(0.162)***	(0.014)***	(0.151)***	(0.163)***	(0.014)***
age^2	-0.022	-0.012	-0.001	-0.022	-0.012	-0.001
	(0.001)***	(0.002)***	(0.000)***	(0.001)***	(0.002)***	(0.000)***
Constant	3.110	16.064		6.884	18.393	
	(5.426)	(4.708)***		(4.807)	(4.763)***	
Observations	829502	515589	842056	810698	507171	823200
No. of clusters/countries	14	14	14	13	13	13
R-squared	0.26	0.04	0.25	0.27	0.04	0.25

Robust standard errors in parentheses (clustered by country), + Marginal effects reported (evaluated at the mean of the independent variables), * significant at 10%; ** significant at 5%; *** significant at 1%

An interesting result appears in Table 8. Although childcare subsidies are positively correlated with participation in the labour market across countries, they are also associated with shorter work weeks for those who are employed. Therefore, these two effects cancel each other out, and the total effect of childcare subsidies on working hours in column (4), after controlling for other institutional factors, is close to zero and statistically insignificant. In the preferred specification of Table 8 (columns (4) – (6)) the effects of family cash benefits are statistically insignificant. While the positive effect of childcare subsidies on participation in column (6) is supportive of the theory, the negative effect on working hours of employed people in column (5) is inconsistent with the theory. One possible explanation is, of course that I am not controlling for other country level characteristics. It is possible, for example that in countries with higher childcare subsidies there is a higher incidence of part-time work, for some unobservable reasons, and that the measure of childcare subsidies is simply picking up the effect of this. Another possible explanation is that in countries with higher childcare subsidies people tend to work shorter hours due to cultural factors. I will explore this in more detail in Section 7.

To further examine possible explanations for the result in Table 7 column (3), where I find no effects of family policies on working hours, in Table 9 I report the results of an analogous analysis, but separately for men and women. Intuitively, one would expect that family support policies would have their effects mostly concentrated on women, since women still bear most of the childcare responsibilities within the household. For both groups the total effects of childcare subsidies on working hours are statistically insignificant, as reported in columns (1) and (4). In both cases, again, the effects on hours worked of employed people are negative and statistically significant for men (columns (2) and (5)), and the effects on participation for both groups are positive and statistically significant (columns (3) and (6)). The effects of family cash benefits are mostly insignificant, except in column (6), where they seem to have significantly positive effect on participation of women. The tax wedge has significantly negative effects in the regressions for men, but it is too imprecisely measured in the case of women[17].

From the results so far I tentatively conclude that there is weak or no evidence that childcare subsidies have an effect on working hours in the economy as predicted by the simple model in Section 2. In the next section I turn my attention to comparisons of people with and without children.

5.4 Comparisons of people with and without children

It is good to first get an idea of the difference in working hours and labour market participation between people with children and people without children. I split people into two groups: the first group includes people living in households with at least one child of age 0-15 present and the second group includes all the rest[18]. Table 10 reports differences across countries between parents and nonparents in their working hours and employment rates. It is based on the sample of age 16-55. As seen from the table, parents on average work 0.4 hours more per week than nonparents, however, there is considerable variation in the differences across countries. In the UK, for example, parents work 5.9 hours less than nonparents, while in Sweden they work 6.7 hours *more*. Similar conclusions can be inferred from the data on the employment rates, calculated from the ECHP and the US CPS data. Again, on average, the

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[17] In order to account for a large number of zeroes among the observed hours of work I repeated the above analysis using a Tobit specification. The results were similar.

[18] One could argue that the most relevant group for analysing the effects of family policies would be parents of children under the age of 6. However, in many countries, families with older children are also beneficiaries of favourable tax breaks. The same goes for family allowances. In any case, later in the analysis I also split parents into two sub-groups: parents with young children and parents with older children.

employment rate is 2.2 % points higher for parents. However, in Luxembourg and the UK it is 8.9 and 8.2 % points lower, respectively, whereas in Sweden it is 17.0 % points higher.

Table 9: Participation, working hours of employed people and family policy: regressions for males and females on the sample of age 16 and more (1998-2001)

Gender group:	Males			Females		
Dependent variable:	Working hours per week		Employed dummy	Working hours per week		Employed dummy
Sample:	all	employed	all	all	employed	all
	OLS		Probit+	OLS		Probit+
	(1)	(2)	(3)	(4)	(5)	(6)
childcare subsidies - % GDP	-0.429	-2.439	0.101	1.228	-2.114	0.150
	(0.670)	(0.590)***	(0.013)***	(1.461)	(1.622)	(0.049)***
family cash benefits - % GDP	-0.176	0.737	0.004	-0.021	-0.932	0.044
	(0.298)	(0.419)	(0.008)	(0.891)	(0.998)	(0.023)*
tax wedge	-31.559	-14.615	-0.434	-26.449	6.131	-0.591
	(6.366)***	(7.737)*	(0.148)***	(17.501)	(18.347)	(0.516)
unemployment rate	-0.265	0.192	-0.013	-0.329	0.253	-0.016
	(0.109)**	(0.111)	(0.002)***	(0.219)	(0.285)	(0.005)***
union density	0.069	0.057	-0.001	0.092	0.040	-0.000
	(0.015)***	(0.020)**	(0.000)***	(0.045)*	(0.052)	(0.001)
employment protection	0.673	1.672	-0.009	-0.320	-0.102	-0.011
	(1.520)	(1.508)	(0.044)	(4.286)	(4.732)	(0.094)
age	2.621	1.349	0.074	1.084	0.697	0.046
	(0.123)***	(0.178)***	(0.013)***	(0.200)***	(0.203)***	(0.014)***
age^2	-0.032	-0.015	-0.001	-0.015	-0.009	-0.001
	(0.001)***	(0.002)***	(0.000)***	(0.002)***	(0.002)***	(0.000)***
Constant	-2.408	18.257		13.401	19.425	
	(3.663)	(4.646)***		(6.581)*	(5.877)***	
Observations	387639	274370	393591	423059	232801	429609
No. of clusters/countries	13	13	13	13	13	13
R-squared	0.34	0.07	0.32	0.23	0.03	0.22

Robust standard errors in parentheses (clustered by country), + Marginal effects reported (evaluated at the mean of the independent variables), * significant at 10%; ** significant at 5%; *** significant at 1%

I now turn to the regression analysis. Recall from Section 2 that we expect a positive effect of childcare subsidies on the working hours of both parents and nonparents, and that these effects are expected to be stronger for parents. Similarly, cash benefits should reduce the working hours of parents and increase the working hours of nonparents. In absolute value the effects of cash benefits are expected to be stronger for parents. I report the results in Table 11. In order to capture possible differences in the effects of family policies on parents and nonparents I include dummy variables (and corresponding interactions) indicating whether children are present in the household. In this way I recognize the potential for heterogeneity in the effects of public policies across groups of individuals; I allow for heterogeneity in the effects of childcare subsidies, cash benefits and the tax wedge. In columns (1) and (2) of the Table 11 I split people into two groups: the first group includes everyone with children of age 0-15 in the household, and the second group includes everyone else. I report two different specifications, one with policy variables only (family policies and taxes) and one with other institutional controls included (unemployment rate and labour market rigidities).

Table 10: Comparisons of people with and without children: working hours and employment rates for the sample of individuals aged 16-55 (year 2000)

	Working hours			% employed		
Country	Children (0-15) present	No children (0-15) present	Difference	Children (0-15) present	No children (0-15) present	Difference
Austria	31.8	32.9	-1.1	80.6	80.4	0.2
Belgium	30.2	26.4	3.8	77.2	65.8	11.4
Denmark	33.0	31.3	1.7	86.5	85.4	1.2
Finland	33.2	29.4	3.8	80.9	73.2	7.6
France	26.4	24.2	2.3	71.4	65.6	5.8
Germany	28.1	31.6	-3.5	75.6	80.5	-4.9
Greece	29.4	25.2	4.2	66.8	59.1	7.7
Ireland	24.9	27.5	-2.6	66.9	71.6	-4.7
Italy	25.4	23.0	2.4	65.1	58.1	7.0
Luxembourg	27.8	32.2	-4.4	75.9	84.8	-8.9
Netherlands	25.3	28.4	-3.1	79.6	81.4	-1.8
Portugal	33.8	30.2	3.7	79.8	72.1	7.7
Spain	24.6	23.8	0.8	58.7	57.3	1.4
Sweden	33.9	27.1	6.7	85.7	68.7	17.0
United Kingdom	28.2	34.1	-5.9	75.0	83.2	-8.2
United States	31.7	33.7	-2.1	81.0	84.4	-3.4
Average	**29.2**	**28.8**	**0.4**	**75.4**	**73.2**	**2.2**

Data Sources: ECHP and IPUMS CPS (King et. al. (2004)).

The results in Table 11 do not support the theoretical predictions. Let me discuss the preferred specification in column (2). The effects of childcare subsidies for both parents and nonparents are insignificant. If anything, the effects are stronger for nonparents (as in column (1)). The effects of cash benefits are also insignificant. The effects of the tax wedge are negative, but statistically significant only for nonparents.

Intuitively, we would expect the results to be strongest for parents with small children. Therefore, in columns (3) and (4) of Table 11 I further split parents into two groups: parents with small children (children of age 0-6 in household) and parents with older children (children of age 7-15 in household and no younger children). The group with no children is the same as in columns (1) and (2). The results are again inconsistent with the theory. The effects of family policies in column (4) are small and statistically insignificant for all groups.

Table 11: Working hours and family policy for people with and without children: regressions on the sample of individuals aged 16 and more (1998-2001)

Dependent variable: Working hours per week				
	(1)	(2)	(3)	(4)
Effects of childcare subsidies				
people with children 0-15	2.252	0.142		
	(1.437)	(1.169)		
people with children 0-6			1.724	-0.371
			(1.389)	(1.008)
people with children 7-15			2.753	0.673
			(1.725)	(1.580)
people without children 0-15	2.827	0.446	2.827	0.451
	(0.794)***	(0.917)	(0.794)***	(0.914)
Effects of family cash benefits				
people with children 0-15	0.749	0.025		
	(1.208)	(0.831)		
people with children 0-6			0.387	-0.206
			(1.252)	(0.880)
people with children 7-15			1.025	0.204
			(1.326)	(1.036)
people without children 0-15	0.590	-0.236	0.590	-0.234
	(0.763)	(0.538)	(0.763)	(0.538)
Effects of tax wedge				
people with children 0-15	-9.361	-12.216		
	(12.684)	(12.846)		
people with children 0-6			-6.196	-9.433
			(12.128)	(12.037)
people with children 7-15			-12.288	-15.068
			(14.611)	(15.205)
people without children 0-15	-31.326	-35.386	-31.326	-35.464
	(9.762)***	(10.631)***	(9.762)***	(10.562)***
children 0-15 in household dummy	-2.085	0.944		
	(5.734)	(5.310)		
children 0-6 in household dummy			-2.346	0.547
			(5.253)	(4.919)
children 7-15 in household dummy			-1.735	1.455
			(6.669)	(6.271)
no 0-15 children in household dummy	5.595	9.659	5.590	9.675
	(5.458)	(4.690)*	(5.448)	(4.676)*
age	1.752	1.751	1.752	1.750
	(0.136)***	(0.142)***	(0.136)***	(0.142)***
age^2	-0.022	-0.022	-0.022	-0.022
	(0.001)***	(0.001)***	(0.001)***	(0.001)***
unemployment rate		-0.302		-0.300
		(0.150)*		(0.150)*
union density		0.083		0.083
		(0.026)***		(0.026)***
employment protection		0.004		0.012
		(2.824)		(2.810)
Observations	829040	810236	829040	810236
No. of clusters/countries	14	13	14	13
R-squared	0.64	0.64	0.64	0.64

Robust standard errors in parentheses (clustered by country), * significant at 10%; ** significant at 5%; *** significant at 1%

Finally, I report results separately for men and women. I also consider the participation decision and working hours' decision separately and I include all other institutional controls. The results are reported in Table 12. Looking at total effects in columns (1) for men, there is evidence of a negative and significant effect of childcare subsidies on the working hours of male parents and a close to zero effect for nonparents. In both cases, participation in the labour market increases with the childcare subsidy (column (3)), whereas hours conditional on working decrease with childcare subsidies (column (2)). Comparing male parents to male nonparents, the difference in the total effects goes in completely wrong direction, as we would expect the effect for parents to be more strongly positive. The positive effect of childcare subsidies on participation in column (3) is encouraging, but this effect is smaller for male parents than for male nonparents. The total effects of cash subsidies for men in column (1) go in the wrong direction; we would expect negative effects for parents and positive effects for nonparents but the effects go in the opposite direction.

Let me now turn to the effects on women. Here the direction of effects is more consistent with the theory, but the coefficients are very imprecisely measured, hence they are in most cases statistically insignificant. The total effects of childcare subsidies in column (4) are positive and stronger for parents; however, they are insignificant. For women, too, there is a positive effect on participation and negative effect on hours worked conditional on working. The effects of childcare subsidies on participation (column (6)) are positive and significant, with no real differences between parents and nonparents. The total effects of cash benefits for women (column (4)) are insignificant, but the effects at least go in the direction expected in theory: there is a negative effect on the working hours of female parents.

To sum up, there is no evidence to support the idea that family policies can help explain differences in working hours across countries. In aggregate, the effects do not show up empirically. There is evidence that childcare subsidies increase participation in the labour market, yet in countries with higher childcare subsidies, people who work tend to work fewer hours. When comparing parents and nonparents, the expected differences in effects and magnitudes also do not show up. The effects on parents do not seem stronger and in many cases go in the wrong direction. If anything, there is some evidence that the effects go in the expected direction for females. However, these effects are very imprecisely measured and are cancelled out by the effects on men. Therefore, they are not enough to be able to explain *aggregate* differences in working hours across countries.

Table 12: Participation, working hours of employed people and family policy for people with and without children: regressions for males and females on the sample of age over 16 (1998-2001)

Gender group:	Males			Females		
Dependent variable	Working hours per week		Employed dummy	Working hours per week		Employed dummy
Sample:	all		employed	all		employed
	OLS		Probit+	OLS		Probit+
	(1)	(2)	(3)	(4)	(5)	(6)
Effects of childcare subsidies						
people with children 0-15	-1.748	-2.170	0.069	2.178	-1.491	0.147
	(0.517)***	(0.609)***	(0.016)***	(1.896)	(2.062)	(0.032)***
people without children 0-15	0.130	-2.617	0.109	0.680	-2.629	0.152
	(0.658)	(0.573)***	(0.018)***	(1.574)	(1.512)	(0.065)**
Effects of family cash benefits						
people with children 0-15	0.727	1.053	0.009	-0.500	-2.123	0.031
	(0.287)**	(0.354)**	(0.014)	(1.466)	(1.486)	(0.029)
people without children 0-15	-0.601	0.529	0.001	0.053	-0.291	0.049
	(0.316)*	(0.466)	(0.009)	(0.840)	(0.827)	(0.027)*
Effects of tax wedge						
people with children 0-15	-15.825	-18.051	-0.025	-11.074	12.227	-0.280
	(6.052)**	(7.541)**	(0.175)	(22.287)	(21.392)	(0.544)
people without children 0-15	-36.347	-12.639	-0.509	-32.445	4.071	-0.736
	(6.681)***	(7.876)	(0.166)***	(17.284)*	(18.936)	(0.543)
children 0-15 in household	-4.828	19.610	-0.078	6.300	13.450	-0.199
dummy	(2.950)	(4.732)***	(0.095)	(8.439)	(8.310)	(0.121)
no 0-15 children in household	-0.284	17.967		17.384	18.498	
dummy	(3.810)	(4.668)***		(6.291)**	(5.989)***	
unemployment rate	-0.269	0.194	-0.014	-0.340	0.226	-0.016
	(0.101)**	(0.110)	(0.002)***	(0.224)	(0.293)	(0.005)***
union density	0.069	0.057	-0.001	0.091	0.039	-0.001
	(0.014)***	(0.020)**	(0.000)***	(0.047)*	(0.055)	(0.001)
employment protection	0.438	1.679	-0.015	-0.507	-0.197	-0.014
	(1.431)	(1.491)	(0.041)	(4.423)	(4.998)	(0.097)
age	2.523	1.329	0.071	1.084	0.888	0.046
	(0.113)***	(0.181)***	(0.012)***	(0.195)***	(0.191)***	(0.014)***
age^2	-0.030	-0.015	-0.001	-0.015	-0.011	-0.001
	(0.001)***	(0.002)***	(0.000)***	(0.002)***	(0.002)***	(0.000)***
Observations	387446	274219	393390	422790	232594	429315
R-squared	0.75	0.94		0.56	0.90	

Robust standard errors in parentheses (clustered by countries). There are 13 clusters (countries) in each regression.
+ Marginal effects reported (evaluated at the mean of the independent variables), in case of dummy variables effects are computed for discrete change of dummy variable from 0 to 1, * significant at 10%; ** significant at 5%; *** significant at 1%.

6 Effects of family-support policies on childcare at home

In order to further explore the effects of family support policies on hours worked, this section examines a parallel prediction of the model that can be tested in the data: that if a mother is encouraged to take up a job, she will likely reduce her time spent in childcare. This happens either because the child is put in some form of formal external childcare institution or the child is taken care of informally, by grandparents for example. Hence, in theory, a childcare subsidy should reduce the time spent in childcare at home for parents. On the other hand, cash benefits should *increase* the time spent in childcare of parents by increasing the non-earned income of parents and thus encouraging them to stay at home. For people without children the effects on childcare are slightly different. Both childcare subsidies and cash benefits are expected to decrease their time spent in childcare at home. This happens via the government budget constraint (13): by increasing support to families, the government reduces the amount of money it awards as a lump sum subsidy, and thus nonparents are induced to work more and spend less time in other activities.

Information on the time spent in childcare is based on the ECHP variable pr007: Number of hours (per week) spent looking after children. This variable is based on two related survey questions: "Do your present daily activities include looking after children, whether your own or other, without pay?" and if the answer to this question is yes, "Roughly how many hours per week do you spend looking after children?" This variable is reported for those who did spend some time looking after children. For the rest (if applicable) I assume they spend 0 hours looking after children. This variable is not available for the US and also not for Germany, Sweden, the UK and Luxembourg.

In Table 13 I show a comparison between childcare at home for people with and without a child aged 0-15 in the household. Countries exhibit a very large variation in the time spent in childcare at home and there is also large variation in the differences in childcare time between the two types of individuals. People with children spend 31.4 hours looking after children in Denmark, but only 13.9 hours in Portugal. On the other hand, people without children spend 0.7 hours per week looking after children in Denmark, France or Spain and 3.22 hours in Italy. On average parents spend 20.2 hours more in childcare at home than nonparents. This difference is highest in Denmark (30.7), Netherlands (28.9) and Ireland (27.9) and lowest in Portugal (13.0), Italy (15.7), Greece (16.0) and France (16.1) In general, south European countries such as Portugal, Greece and Italy tend to have low differences, indicating that there may be a lot of informal care for children going on, i.e. by grandparents.

Table 13: Time spent in childcare at home for people with and without children, sample of age 16 and more (year 2000)

Country	Children (0-15) present	No children (0-15) present	Difference
Austria	18.4	0.9	17.5
Belgium	21.9	1.9	20.1
Denmark	31.4	0.7	30.7
Finland	19.4	0.8	18.6
France	16.8	0.7	16.1
Greece	17.7	1.7	16.0
Ireland	29.0	1.1	27.9
Italy	18.9	3.2	15.7
Netherlands	31.1	2.3	28.9
Portugal	13.9	0.9	13.0
Spain	18.3	0.7	17.5
Average	21.5	1.4	20.2

Data Sources: ECHP. Time spent in childcare is based on ECHP variable pr007: Number of hours (per week) spent looking after children. This variable is based on two related survey questions: "Do your present daily activities include looking after children, whether your own or other, without pay?" and if the answer to this question is yes, "Roughly how many hours per week do you spend looking after children?" This variable is reported for those who did spend some time looking after children. For the rest (if applicable) I assumed they spend 0 hours looking after children.

In Table 14 I report regression results from regressing the hours looking after children on the policy variables. Because data on childcare is not available for quite a few countries, only 10 countries are left for this analysis. Therefore, the results of this section cannot be directly comparable to the results from the previous section, where I analysed the effects of policies on working hours. Columns (1) and (2) of Table 14 report results from regressions on the whole sample with institutional controls either excluded or included, respectively. In aggregate, childcare subsidies are expected to *decrease* time in childcare at home, whereas cash benefits are expected to *increase* time in childcare. The empirical results, however, are not consistent with these predictions. In the specification with other controls included, column (2), the effect of childcare subsidies is statistically insignificant, but positive. The effect of cash benefits on the other hand is significantly negative. In columns (3) and (4) I do analogous analysis, but allowing for differential effects between parents and nonparents. Recall that the effect of childcare subsidies is expected to be more negative for parents, whereas the effect of family cash benefits is supposed to be positive for parents and negative for nonparents. This is not the case in my results. The effects of childcare subsidies on parents are in fact positive and significant, whereas for nonparents they are negative and significant. This means that in countries with higher childcare subsidies parents appear to take more care of children at home! The only statistically significant effect that goes in the right direction is the effect of cash subsidies on nonparents in column (4). Finally, in columns (5) and (6) I split the sample by gender, and again results are inconsistent with the theory for both males and females.

Table 14: Childcare at home and family policy for people with and without children: regressions on the sample of age over 16 (1998-2001)

Dependent variable: Hours per week looking after children						
Sample:	all	all	all	all	males	females
	(1)	(2)	(3)	(4)	(5)	(6)
Effects of childcare subsidies						
aggregate	2.189	0.719				
	(0.922)**	(0.543)				
people with children 0-15			9.523	7.705	9.533	5.819
			(2.324)***	(1.719)***	(1.036)***	(2.501)**
people without children 0-15			-1.018	-2.790	-1.465	-4.044
			(0.498)*	(0.708)***	(0.567)**	(0.910)***
Effects of family cash benefits						
aggregate	-0.434	-1.022				
	(0.645)	(0.449)*				
people with children 0-15			-0.254	-1.034	-0.861	-1.339
			(1.618)	(1.290)	(0.879)	(1.878)
people without children 0-15			-0.492	-1.118	-0.493	-1.650
			(0.414)	(0.449)**	(0.386)	(0.587)**
Effects of tax wedge						
aggregate	-11.019	-5.962				
	(13.376)	(5.060)				
people with children 0-15			-54.434	-50.417	-33.063	-63.657
			(38.202)	(27.113)	(18.205)	(37.790)
people without children 0-15			9.941	17.262	3.931	27.329
			(8.893)	(6.808)**	(4.459)	(10.927)**
children 0-15 in household dummy			29.152	40.080	18.175	58.554
			(16.264)	(12.735)**	(8.112)*	(18.421)**
no 0-15 children in household dummy			-10.237	-1.278	0.300	-3.061
			(4.570)*	(4.161)	(2.122)	(6.865)
unemployment rate		-0.052		-0.054	-0.053	-0.026
		(0.057)		(0.070)	(0.038)	(0.106)
union density		-0.026		-0.030	0.006	-0.062
		(0.019)		(0.016)*	(0.007)	(0.026)**
employment protection		-8.152		-10.034	-5.285	-14.917
		(1.026)***		(1.373)***	(0.880)***	(2.032)***
age	0.639	0.636	0.431	0.424	0.236	0.656
	(0.107)***	(0.112)***	(0.097)***	(0.102)***	(0.041)***	(0.166)***
age^2	-0.008	-0.008	-0.004	-0.004	-0.002	-0.007
	(0.001)***	(0.001)***	(0.001)***	(0.001)***	(0.000)***	(0.002)***
Constant	2.075	10.190				
	(6.346)	(4.004)**				
Observations	333447	314620	333385	314558	151885	162673
No. of clusters/countries	10	9	10	9	9	9
R-squared	0.06	0.06	0.38	0.39	0.32	0.52

Robust standard errors in parentheses (clustered by countries); * significant at 10%; ** significant at 5%; *** significant at 1%

In order to better understand what is behind these results, I next split the sample by age into two groups, people aged 16-55 and people who are older than 55 years. I choose the 55 years cut-off because this is approximately the average age at which people become grandparents. I report the results in Table 15. In columns (1) and (2) I report the results for the sample of those aged 16-55. Similarly as before, there is a puzzling positive effect of childcare subsidies on time in childcare activities at home and a negative effect of cash benefits, column (1); both these effects are inconsistent with the predictions of the theory. Comparing parents and nonparents, the effects of childcare subsidies are strongly positive for parents, which is also puzzling. Moving to column (3), where I show results from the analysis on people of age 56 or more, we can see significant and negative effects of both types of family policies on time spent in childcare at home. There seems to be some indication that family policies do not have the expected effects on the prime-age group, or on parents, but there seem to be strong and significant effects on old people, for example grandparents. One potential explanation for this would be that if family policies are not relieving the childcare burden of parents, perhaps they are merely relieving the childcare burden of grandparents. Therefore, because grandparents are often not active in the labour market, the effects on working hours cannot be large. However, there may be certain cultural or institutional unobservables that are correlated with family policy variables and are thus contaminating the results. I explore this question further in the next section.

7 Controlling for cultural factors

In a cross-country analysis there is always a risk of unobservable country-level variables that are correlated with observed regressors contaminating the results. My analysis is very much prone to this kind of bias. I have discussed above that it is possible that cultural aspects affect my results and bias the coefficients on family policy variables. For example, in Section 5.3 I report that conditional on working, childcare subsidies have a negative effect on working hours. However, it could as well be the case that in countries with high public childcare support, part-time work is more prevalent. Furthermore, I also find that childcare subsidies have a positive effect on time spent in childcare at home by parents. This result, however could be caused by the fact that in countries with lower childcare subsidies, such as southern Europe, parents take less time to care for their children because using informal care by other family members is more readily available and more acceptable. The reported effect could thus appear merely due to cultural differences across countries that are correlated with the measure of childcare subsidies.

Table 15: Childcare at home and family policy for people with and without children: regressions on the samples of individuals aged 16-55 and over 55 (1998-2001)

Dependent variable: Hours per week looking after children			
Age group:	16-55		>55
	(1)	(2)	(3)
Effects of childcare subsidies			
aggregate	1.942		-1.875
	(0.584)**		(0.327)***
people with children 0-15		6.581	
		(1.373)***	
people without children 0-15		-2.769	
		(0.896)**	
Effects of family cash benefits			
aggregate	-0.953		-1.088
	(0.479)*		(0.263)***
people with children 0-15		-1.381	
		(1.080)	
people without children 0-15		-0.888	
		(0.616)	
Effects of tax wedge			
aggregate	-22.950		23.768
	(5.586)***		(2.818)***
people with children 0-15		-54.580	
		(21.538)**	
people without children 0-15		7.074	
		(8.207)	
children 0-15 in household dummy		15.526	
		(9.572)	
no 0-15 children in household dummy		-24.102	
		(6.934)***	
unemployment rate	0.019	0.008	-0.123
	(0.064)	(0.072)	(0.031)***
union density	-0.012	-0.020	-0.019
	(0.021)	(0.016)	(0.011)
employment protection	-9.819	-12.289	-1.193
	(1.087)***	(1.358)***	(0.618)*
age	4.062	2.215	-0.083
	(0.404)***	(0.217)***	(0.060)
age^2	-0.056	-0.030	-0.000
	(0.006)***	(0.003)***	(0.000)
Constant	-38.294		3.443
	(6.266)***		(2.353)
Observations	217131	217077	97489
R-squared	0.10	0.42	0.03

Robust standard errors in parentheses (clustered by countries). There are 9 clusters (countries) in each regression. * significant at 10%; ** significant at 5%; *** significant at 1%

In order to assess the impact of this I include into the analysis two more variables. The first measures the prevalence of part time work in a country, and the second measures the average number of adults in a household. Both variables are calculated from the information available in the ECHP and the US CPS. In Table 16 I show these two variables across countries for the year 2000. The share of part-time workers measures the percentage of employed people who are employed part-time. The share of part-time workers is partly determined by institutional arrangements in the labour market, and partly by preferences of consumers/households. As evident from Table 16, 13% of workers are on average employed part-time[19], but differences across countries are very large. In the Netherlands, for example 31.2% employed people work part-time. This is followed by Ireland (19.1%) and the US (17.5%). At the bottom end of the spectrum are southern European countries, Greece (5.5%), Italy (6.9%) and Portugal (7.8%).

Table 16: Measures of cultural factors across countries: Incidence of part-time work and the average number of adults in the household (2000)

Country	% of part-time workers	average number of adults in household (>18)
Austria	13.2	2.0
Belgium	15.5	2.0
Denmark	15.3	1.9
Finland	9.7	1.7
France	8.9	2.0
Germany	16.9	1.9
Greece	5.5	2.3
Ireland	19.1	2.3
Italy	6.9	2.3
Luxembourg	na	1.9
Netherlands	31.2	1.8
Portugal	7.8	2.5
Spain	8.9	2.6
Sweden	9.0	1.6
UK	10.0	1.8
US	17.5	1.9
Average	**13.0**	**2.0**

Data Sources: ECHP and IPUMS CPS (King et al. (2004)).

Another variable measures the average number of adults aged 18 or more in a household. With this measure I seek to capture cultural differences across countries regarding family size, living arrangements and the proximity of potential informal care for children. The unweighted average of the number of adults in the household across countries is equal to 2.0. As expected, the biggest size of households is observed in southern Europe: Spain (2.6), Portugal (2.5), Greece and Italy (both 2.3). Ireland also has a high average number of adults (2.3) in a household. At the lower end of the scale are countries such as Sweden (1.6), Finland (1.7), Netherlands (1.8) and the UK (1.8).

[19] Based on unweighted cross-country average.

Table 17: Impact of cultural factors – the average number of adults in the household and the incidence of part-time work: Working hours regressions on the sample of age over 16 (1998-2001)

Dependent variable:	Working hours per week		Employed dummy	Working hours per week		Employed dummy
Sample:	all	employed	all	all	employed	all
	OLS		Probit+	OLS		Probit+
	(1)	(2)	(3)	(4)	(5)	(6)
Effects of childcare subsidies						
aggregate	1.220	-1.142	0.135			
	(0.405)**	(0.450)**	(0.018)***			
people with children 0-15				0.975	-0.809	0.117
				(0.764)	(0.600)	(0.031)***
people without children 0-15				1.208	-1.337	0.141
				(0.405)**	(0.483)**	(0.023)***
Effects of family cash benefits						
aggregate	-0.251	0.392	0.004			
	(0.392)	(0.376)	(0.011)			
people with children 0-15				-0.142	0.113	-0.003
				(0.701)	(0.513)	(0.021)
people without children 0-15				-0.442	0.574	0.004
				(0.358)	(0.345)	(0.011)
Effects of tax wedge						
aggregate	-45.602	-14.183	-1.040			
	(10.289)***	(9.461)	(0.278)***			
people with children 0-15				-28.710	-12.565	-0.630
				(13.724)*	(11.210)	(0.331)*
people without children 0-15				-52.186	-14.857	-1.210
				(9.813)***	(9.546)	(0.290)***
share of part-time workers	-30.326	-28.809	-0.758	-30.512	-29.190	-0.767
	(7.026)***	(4.790)***	(0.208)***	(7.229)***	(5.030)***	(0.214)***
average number of adults (>18) in household	-2.932	1.270	-0.180	-3.187	1.439	-0.186
	(1.148)**	(1.003)	(0.038)***	(1.197)**	(1.049)	(0.038)***
children 0-15 in household dummy				17.493	20.906	-0.204
				(6.875)**	(6.523)***	(0.097)**
no 0-15 children in household dummy				26.539	22.977	
				(5.659)***	(5.378)***	
unemployment rate	-0.410	0.041	-0.015	-0.412	0.029	-0.015
	(0.052)***	(0.048)	(0.001)***	(0.054)***	(0.050)	(0.002)***
union density	0.082	0.039	-0.001	0.082	0.038	-0.001
	(0.013)***	(0.012)***	(0.000)**	(0.014)***	(0.013)***	(0.000)*
employment protection	2.181	0.698	0.092	2.127	0.625	0.090
	(1.507)	(1.218)	(0.045)**	(1.579)	(1.293)	(0.046)*
age	1.785	1.088	0.060	1.759	1.168	0.060
	(0.149)***	(0.166)***	(0.014)***	(0.140)***	(0.178)***	(0.014)***
age^2	-0.022	-0.013	-0.001	-0.022	-0.014	-0.001
	(0.001)***	(0.002)***	(0.000)***	(0.001)***	(0.002)***	(0.000)***
Constant	23.001	23.696				
	(5.800)***	(5.387)***				
Observations	810698	507171	823200	810236	506813	822705
R-squared	0.27	0.06	0.26	0.65	0.91	0.26

Robust standard errors in parentheses (clustered by countries). There are 13 clusters (countries) in each regression. + Marginal effects reported (evaluated at the mean of the independent variables), in case of dummy variables effects are computed for discrete change of dummy variable from 0 to 1; z values in parentheses (test of the underlying coefficient being 0), * significant at 10%; ** significant at 5%; *** significant at 1%

The first set of regression results is reported in Table 17. I repeat the analysis from before with the addition of the two new measures of cultural factors. I turn first to the effects of the variables measuring cultural factors. They seem to have significant effects in most of the specifications. The incidence of part time work enters with a negative and significant coefficient in all six specifications, as expected. Similarly, the average number of adults has a negative and significant effect in 4 out of 6 cases, also in line with the intuition. In countries with larger families, people tend to work less on average, mainly due to lower participation in the labour market.

Columns (1) - (3) of Table 17 are analogous to the columns (4) - (6) in Table 8. They analyse the aggregate effects of family policies on working hours in the economy, decomposed into the participation and hours of work decision. First, with culture included in the regression, there is a positive and significant total effect of childcare subsidies on working hours (column (1))). In column (2), despite including the measure of the incidence of part-time work into the regression, there is still a significantly negative, but smaller than before, effect of childcare subsidies on working hours conditional on working. Hence, this effect seems to be quite robust. In column (3) the effect of childcare subsidies on participation is more or less the same as before. The effects of cash benefits on working hours are insignificant on aggregate. These results tend to be supportive of the theory, and hence indicate the possibility that after controlling for differences in household size and the incidence of part-time work, family policies have a scope to explain differences in working hours across countries.

However, when comparing parents with nonparents, the differences remain puzzling and contradict the theory. Columns (4) - (6) in Table 17 show the results of this exercise. The effects of childcare subsidies tend to be stronger for nonparents rather than parents, and the effects of cash benefits are all insignificant. The results also remain inconsistent with the theory when analysing the time spent in childcare at home, reported in Table 18 (analogous to results reported in Table 15). One can see that incidence of part-time work and the size of households have strong and significant positive effects on childcare at home; however, this does not prevent the coefficients on family policies from being "wrong". On aggregate, in column (1), childcare subsides still tend to have a positive and significant effect on childcare at home. When comparing parents and nonparents in column (2), the difference in coefficients is still counterintuitive, as the effect on parents is, contrary to expectations, significantly positive. And when analysing the effects on older people, the negative effects of childcare subsidies and cash benefits are reduced after the inclusion of the measures of cultural factors, but remain negative and highly statistically significant. This confirms the idea that family policies relieve the childcare burden of nonparents rather than parents.

Table 18: Impact of cultural factors – the average number of adults in the household and the incidence of part-time work: Childcare at home regressions on the sample of age over 16 (1998-2001)

Dependent variable: Hours per week looking after children			
Age group:	16-55		>55
	(1)	(2)	(3)
Effects of childcare subsidies			
aggregate	2.210 (0.351)***		-1.659 (0.106)***
people with children 0-15		6.429 (1.184)***	
people without children 0-15		-2.941 (0.546)***	
Effects of family cash benefits			
aggregate	-0.317 (0.228)		-0.666 (0.085)***
people with children 0-15		-1.036 (0.884)	
people without children 0-15		-0.603 (0.503)	
Effects of tax wedge			
aggregate	2.943 (7.397)		38.828 (2.657)***
people with children 0-15		-20.228 (17.191)	
people without children 0-15		43.024 (12.158)***	
share of part-time workers	27.399 (4.728)***	36.989 (4.370)***	15.975 (1.371)***
average number of adults (>18) in household	6.237 (1.459)***	5.119 (1.219)***	3.964 (0.426)***
children 0-15 in household dummy		-21.316 (7.281)**	
no 0-15 children in household dummy		-61.655 (9.384)***	
unemployment rate	0.009 (0.042)	0.030 (0.040)	-0.132 (0.013)***
union density	0.004 (0.008)	0.004 (0.008)	-0.011 (0.003)***
employment protection	-7.330 (1.069)***	-7.001 (1.022)***	0.080 (0.259)
age	4.055 (0.405)***	2.198 (0.217)***	-0.092 (0.063)
age^2	-0.056 (0.006)***	-0.030 (0.003)***	-0.000 (0.000)
Constant	-70.510 (9.103)***		-15.690 (3.854)***
Observations	217131	217077	97489
R-squared	0.11	0.42	0.03

Robust standard errors in parentheses (clustered by countries). There are clusters (countries) in each regression.
* significant at 10%; ** significant at 5%; *** significant at 1%

8 Conclusion

In this chapter I test whether fiscal family-support policies can help explain the differences in working-hours across countries. This has been suggested in the literature by Rogerson (2007) and Ragan (2005), but has so far relied on simulations of stylised macroeconomic models with a representative agent. However, the proposed explanations imply some important effects at a more disaggregated level. I therefore focus on differences between people with children and people without children.

I consider two types of public family policies: childcare subsidies and family cash benefits. With a simple theoretical model I show that we would expect the effect of childcare subsidies on aggregate working hours 1) to be positive in general, 2) to be positive for both parents and nonparents, but 3) be stronger for parents. There should be 4) no aggregate effects of family cash benefits on working hours, 5) a negative effect for parents, and 6) a positive effect for nonparents. The effects of family policies are also expected to affect the time in childcare within a household. According to the model outlined in this chapter there is 1) a negative effect of childcare subsidies on aggregate time spent in childcare in general, 2) negative for both parents and nonparents, but 3) stronger for parents. There should be 4) positive aggregate effects of family cash benefits on childcare time, 5) the effects should be positive for parents, and 6) negative for nonparents.

I test these predictions using cross country variation from European Household Panel and US CPS data. In the aggregate setting there is no supportive evidence for the idea that family policies help explain differences in working hours across countries. In preferred specifications with other controls included, the effects are close to zero and insignificant. In countries with higher childcare subsidies, participation in the labour force is indeed higher, but working hours *conditional on working* are lower, bringing the aggregate effect of policies to zero. When comparing the effects of parents and nonparents, the differences in the coefficients are counterintuitive. There is an indication that results go in the right direction for females, but are cancelled out by the counter effects on males; hence, on the whole, family policies do not have the expected effect on working hours.

Furthermore, in regressions with time spent in childcare at home on the left hand side, the effects of policy variables contradict the theory. Childcare subsidies, for example, seem to increase the time in childcare at home and this effect seems to be, counter-intuitively, strongest for parents. An analysis performed on the sample of people older than 55 years, for which family policy should not matter much, seems to indicate that childcare subsidies tend to reduce the time in childcare for this group. One possible explanation for this is that public childcare support actually relieves grandparents from childcare, rather than parents. But, because older people do not participate in the labour market very much, family support cannot have a strong impact on working hours. Finally, I incorporate into the analysis two measures of cultural factors: the average number of adults in the household and the incidence of part-time work in the country. After controlling for these two variables, the aggregate effects of childcare subsidies on working hours become positive and significant; however, the effects on parents are still weaker than the effects on nonparents and the results from the analysis on the time in childcare spent at home still contradict the theory.

The lack of empirical evidence in support of the idea that family policies can help explain the differences in working hours across countries is striking. The family policy story can perhaps be used narrowly to explain the differences between Sweden and the US, but it does not bear the inclusion of a greater set of countries and a greater set of controls. Nor does it bear the separation of effects for parents and nonparents. I therefore conclude that the family policy story contributes little to the explanation of the differences in working hours across countries.

However, it is important to note that in practice, family policies are not implemented solely to affect labour supply. Instead, they may be directed at other objectives such as child protection, child development and education, parental health, gender equity and fertility. Indeed, some of these, e.g. fertility and gender equity, can affect labour supply in their own way, possibly masking the effects of child subsidies and cash benefits. Due to data limitations they have not been included in the analysis above, but they remain an important priority for future research.

The objective of this chapter is to test whether the predictions of macroeconomists trying to explain variation in working hours across countries are correct, and this was tested using individual level data. Therefore, the model presented here is more of a macro type, is gender neutral and includes only some minor variation across individuals regarding preferences and public policies. It is possible that very simple assumptions in such macro models give rise to effects that would not appear in more involved models. As exposed for example in Ermisch (1989), in models that take into account decisions within a household regarding labour supply, fertility and childcare, the effects of childcare policies are mostly of an ambiguous sign. This would lead one to conclude that the predictions that macroeconomists talk about are not grounded in microeconomic behaviour, a conclusion that in fact is in line with the empirical conclusions presented in this chapter.

Finally it is important to stress that given the slow variation in policies over time, the analysis is done mainly by exploiting variation across countries. However, with data available only for a relatively small number of countries, the power of the analysis to address this sort of question is arguably low. Hence, with more variation in public policies over time, or when data becomes available for a greater number of countries, future analysis may lead to more robust conclusions.

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Appendix

Here I show some results from the Section 2 analytically.

First order conditions

For completeness I state the first order conditions for maximising utility subject to specified constraints.

Parents FOC (based on equations (1) – (4)):

$$\frac{\partial L}{\partial h_P}: \frac{\alpha(1-\tau)}{(1-\tau)h_P + T + s} = \frac{1-\alpha-\beta}{1-h_P-H_P}$$

$$\frac{\partial L}{\partial H_P}: \frac{\beta}{H_P + g} = \frac{1-\alpha-\beta}{1-h_P-H_P}$$

Nonparents FOC (based on equations (5) – (8)):

$$\frac{\partial L}{\partial h_N}: \frac{\alpha(1-\tau)}{(1-\tau)h_N + T} = \frac{1-\alpha-\gamma}{1-h_N-H_N}$$

$$\frac{\partial L}{\partial H_N}: \frac{\gamma}{H} = \frac{1-\alpha-\gamma}{1-h_N-H_N}$$

The effects of taxes on labour supply

Aggregate labour supply and labour supply of parents unambiguously decrease with taxes:

$$\frac{dh}{d\tau} = -\frac{(1-\alpha^2)}{(1-\alpha\tau)^2} < 0$$

$$\frac{dh_P}{d\tau} = -\frac{(1-\alpha^2)}{(1-\alpha\tau)^2} - s\frac{(1-\alpha)(1-\delta)}{(1-\tau)^2} < 0$$

For the labour supply of nonparents:

$$\frac{dh_N}{d\tau} = -\frac{(1-\alpha^2)}{(1-\alpha\tau)^2} + s\frac{(1-\alpha)\delta}{(1-\tau)^2} < 0, \text{ IFF } \frac{(1+\alpha)}{(1-\alpha\tau)^2} > s\frac{\delta}{(1-\tau)^2}$$

Under plausible conditions, s and δ are much smaller than 1, hence this condition will be satisfied. See "Realistic example" below.

The effects on childcare at home

Childcare from FOC:

$$H_P = \frac{\beta - g(1-\alpha-\beta)}{1-\alpha} - \frac{\beta}{1-\alpha} h_P ,$$

$$H_N = \frac{\gamma}{1-\alpha} - \frac{\gamma}{1-\alpha} h_N .$$

Effects of taxes on childcare:

$$\frac{dH}{d\tau} = \frac{\alpha(\alpha\beta + (1-\delta)\gamma)}{(1-\alpha\tau)^2} + s\frac{(\beta-\gamma)\delta(1-\delta)}{(1-\tau)^2} > 0$$

$$\frac{dH_P}{d\tau} = \frac{\alpha\beta}{(1-\alpha\tau)^2} + s\frac{\beta(1-\delta)}{(1-\tau)^2} > 0$$

And,

$$\frac{dH_N}{d\tau} = \frac{\alpha\gamma}{(1-\alpha\tau)^2} - s\frac{\gamma\delta}{(1-\tau)^2} > 0 \text{, IFF } \frac{\alpha}{(1-\alpha\tau)^2} > s\frac{\delta}{(1-\tau)^2} .$$

The effect on childcare of nonparents is ambiguous, although again, in realistic cases, s should be much smaller than α and the condition for a positive sign would hold.

"Realistic example"

Suppose $\alpha = 0.4, \tau = 0.4$, $\delta = 0.4$ and $g = 0$; then $h \approx 0.2857$. If government gives 10% of the GDP (deliberate overestimate) for family cash benefits then $s = 0.0714$. Then both conditions above hold by a wide margin. One can think of an exhaustive range of "realistic examples" where the above conditions would always hold.

Alternative set up for the government subsidy *g* with the market for childcare

Alternatively, the utility maximisation problem of parents could look like this:

$$\alpha \ln c_P^m + \beta \ln c_P^h + (1-\alpha-\beta) \ln l_P$$

subject to

$$c_P^m + (1-g)C = (1-\tau)h_P + T + s$$

$$c_P^h = (H_P^\rho + C^\rho)^{1/\rho}$$

$$l_P = 1 - h_P - H_P ,$$

where C represents the "amount" of childcare that parents buy on the market and ρ determines the elasticity of substitution between childcare at home and childcare provided on the market.

In this set up, the government pays for a share g of the childcare C that a household buys on the market. The government hence implicitly reduces the price of childcare on the market relative to the self provision of childcare at home.

The utility maximisation problem of nonparent households remains unchanged.

Resulting in FOC for parents:

$$\frac{\partial L}{\partial h_P}: \frac{\alpha(1-\tau)}{(1-\tau)h_P + T - (1-g)C} = \frac{1-\alpha-\beta}{1-h_P-H_P}$$

$$\frac{\partial L}{\partial H_P}: \frac{\beta H_P^{\rho-1}}{H_P^\rho + C^\rho} = \frac{1-\alpha-\beta}{1-h_P-H_P}$$

$$\frac{\partial L}{\partial C}: \frac{\alpha(1-g)}{(1-\tau)h_P + T - (1-g)C} = \frac{\beta C^{\rho-1}}{H_P^\rho + C^\rho}$$

Closed-form solutions are complicated in this kind of set up. It turns out, however, that as long as market- and domestically-provided childcare are close substitutes (the elasticity of substitution is greater than 1), which is plausible, the government subsidy g positively affects working hours h.

Chapter II

Greater Wage Inequality Reduces Average Hours of Work

Urban Sila

Abstract: In this chapter I argue that a mean-preserving spread in offered wages (rising wage inequality) lowers average working hours in the economy. If labour supply is concave in wages, responses to changes in wages are stronger at the bottom of the wage distribution. Hence, a decrease in the working hours of low-paid workers is greater than an increase in working hours of high-paid workers. Furthermore, due to low market opportunities, some of the low-paid workers may leave the labour force and become inactive. Using CPS-MORG data for prime-age men for the 1979-2008 period, I find evidence in support of this explanation. I establish empirically the concavity of the labour supply function and find evidence that after controlling for the average wage, wage inequality has a negative and significant effect on labour supply. This result is robust to various specifications.

1 Introduction

There has been a substantial increase in US wage inequality over the last three decades. The biggest rise in wage inequality among prime-age male workers took place in the 1980s, after which this upward trend somewhat slowed down. The changes in the distribution of wages have attracted a lot of attention from researchers. Two recent contributions by Autor, Katz and Kearney (2008) and Lemieux (2006) undertake extensive empirical analysis of the trends in wage inequality and discuss potential explanations.

Many economists have suggested that shifts in wage inequality can have important implications for the labour supply. Bell and Freeman (2001), for example, argue that forward-looking workers worry about future promotions and pay rises. Therefore, since higher inequality corresponds to higher wage gains, workers work *harder* in industries with more unequal distribution of wages. Similar findings are reported by Kuhn and Lozano (2008) who report that among salaried men, increases in long work hours (above 48 per week) were greatest in industries with the largest increases in wage inequality. Positive effects of inequality on working hours are also reported in Bowles and Park (2005). In their case this is explained by the so-called "Veblen effects". People look up to the richest class in society and seek to match the rich class in their consumption. Hence, the higher the variation in incomes in society, the harder people will want to work in order to imitate the rich class[1].

This branch of literature relies on workers' career concerns and on interpersonal comparisons through which inequality affects working effort of each individual worker. My approach, however, is different. In my work I conceptually follow a more "conventional" labour supply literature such as Juhn, Murphy and Topel (1991), Juhn (1992) and Welch (1997). In this type of setting an individual's labour supply simply traces the wage that that individual (potentially) earns in the labour market. Assuming a positive effect of wages on working hours, if the offered wage of a worker rises, the labour supply of that worker will rise. On the other hand, if the offered wage falls, the worker will decrease her work effort or may drop out of the labour force altogether[2]. If in addition the labour supply function is stable over time

[1] After a careful econometric analysis, Carr (2008) finds no empirical support for the positive causal effect of wage inequality on working hours. He concludes that the positive correlation reported by other researchers is due to endogeneity. A positive effect arises either due to variations in the underlying institutional structure of occupations or due to self-selection into occupations.

[2] Evidence of such behaviour is reported in Juhn et al. (1991) and Juhn (1992). They observe that due to falling wages for male workers with low wages and low skills, an increasing proportion of them are jobless and opting out of the labour force.

then movements in a worker's wage will closely determine movements in his or her labour supply over time. According to such a view of the labour market, it is obvious that changes in the wage inequality should have direct consequences for labour supply.

I argue that a mean-preserving spread in the wage distribution causes average working hours in the economy to *fall*. The reasoning behind this is quite simple. Suppose that the labour supply function is concave in offered wages. When variation in wages increases, keeping the mean wage constant, high-paid workers are getting paid more whereas low-paid workers are getting paid less. Due to the concavity of labour supply, the effects on labour supply at the bottom of the wage distribution will be stronger than at the top of the wage distribution. Therefore, while high-paid workers increase their work hours by little, low-paid workers decrease their labour supply by a lot. Moreover, for some workers offered wages are reduced below their reservation wage and they decide to drop out of the labour force and work zero hours. In practice, active workers usually work a certain amount of hours that is not close to zero. Hence, a decision to leave the labour force usually represents a discreet jump and decreases working hours substantially. By this argument, therefore, higher variation in wages causes working hours in the economy to be lower.

For the results of this chapter I rely heavily on the concavity of the labour supply function. The idea that the labour supply function is concave in offered wages has been around in the literature, at least implicitly, for some time. Much of the established labour-supply literature assumes a concave labour supply, as reported also in Blundell and MaCurdy (1999). In some cases, the concavity of labour supply was reported also explicitly: Juhn et al. (1991) and Juhn and Murphy (1997) find that with rising wages labour supply elasticity falls. Furthermore, in the empirical section of this chapter, I find support for the concavity of labour supply with my data.

Finally, I report empirical evidence in support of the claim that greater wage inequality decreases average hours of work in the economy. I do this using the NBER extracts of the CPS Annual Earnings File (also known as the Merged Outgoing Rotation Groups) for prime-age men for the period 1979-2008. I find evidence that after controlling for the average wage, wage inequality indeed has a negative and significant effect on labour supply. This result is robust to various specifications. According to my results, an increase in wage inequality equivalent to the increase in the US over the 1979-2008 period would cause an approximately 3 % decrease in average hours worked in the economy. During the same period, observed average working hours in the economy have fallen for about 8%. From this it follows that rising inequality can potentially explain 30-40% of the decrease in working hours over time.

The findings of this chapter are important for several reasons. The chapter deals with two central themes in labour economics – labour supply and inequality. To my best knowledge this is the first paper that explicitly relates the conventional static view of labour supply, where labour supply is determined by a worker's own offered wage, with wage inequality. Furthermore, the results of this chapter are also interesting from the point of view of the literature that attempts to explain the differences in working hours across countries (see for example Prescott (2004), Alesina, Glaeser and Sacerdote (2005), Rogerson (2006), Rogerson (2007) and Faggio and Nickell (2007)). In this context wage inequality is another variable that affects working hours in the economy and should be taken into account when comparing working hours across countries. The results are also interesting from the point of view of public policies that affect wage inequality.

The chapter is organised as follows. In section 2 I argue theoretically that the concave labour supply function implies that a mean-preserving spread in wages may result in a fall in average hours worked in the economy. In section 3 I describe the data and show the evolution of wages and wage inequality over time for prime-age men. In section 4 I present the main results of the paper. First, I establish empirically that labour supply is indeed concave in wage. I then test the theoretical prediction that wage inequality, controlling for average wage, has a negative effect on working hours. In section 5 I conclude.

2 Labour supply and wage inequality

2.1 Concavity of labour supply

Theoretically, labour supply is not necessarily concave in wage. However, in much of the empirical labour-supply literature, researchers have implicitly assumed and estimated a concave functional form. In his influential analysis of the sensitivity of an empirical female labour supply model, Mroz (1987) assumes a semi-logarithmic specification. Pencavel (1986) mentions both logarithmic and semi-logarithmic specifications, but reports that the most commonly used specification is simply linear. In their survey of the labour supply literature, Blundell and MaCurdy (1999) report that most often the empirical specifications of the static labour supply model are of two forms:

$$\ln h_i = \beta \ln w_i + \alpha' Z_i + \varepsilon_i$$
$$h_i = \beta \ln w_i + \alpha' Z_i + \varepsilon_i \quad (1)$$

Where h represents hours worked, w is offered wage, Z is a vector of other controls, ε is the error term, α and β are parameters and subscript i stands for an individual. Both specifications in (1) imply that the effect of wages on hours is decreasing with wage, and hence is concave.

The concavity of the labour supply has been reported also more explicitly. Juhn et al. (1991) estimate elasticities of labour supply for men by percentiles of the wage distribution and they find that for workers that are higher up in the wage distribution, cross-sectional estimates of the labour supply elasticities are significantly smaller. Similar findings are reported by Juhn and Murphy (1997); they estimate an employment equation for different groups of married men according to the percentile in the wage distribution and report that employment rate is clearly less responsive for those who are high in the wage distribution. Devereux (2004) also finds that own-wage elasticity is decreasing in the husband's wage.

In the empirical section below I will report evidence that the cross-sectional labour supply curve is concave with respect to wages, consistent with the evidence just cited. For the remainder of this section I will thus assume that the relationship between working hours and wages is concave.

2.2 The effects of wage inequality

Let us assume the following full specification of the labour supply:

$$h_i = 0, \text{ if } w_i \le \overline{w}(X_i), \quad (2)$$

$$h_i = g(w_i) + \alpha' Z_i, \text{ if } w_i > \overline{w}(X_i), \quad (3)$$

where $g(.)$ is a concave function and $\overline{w}(.)$ represents a reservation wage function; X is a vector of personal characteristics which may overlap with the vector Z. Note that function $g(.)$ is defined over the whole range of offered wages, above and below the reservation wage. We can assume, plausibly, that the value of h for w just above the reservation wage $\overline{w}(.)$ is *not* too close to zero. This is plausible because in practice active workers usually work a certain amount of hours that is not very close to zero. I will return to this point below.

Furthermore, assume that offered wages are distributed according to an exogenous distribution defined by the pdf $f(w)$. Assume that individuals differ only by wage and personal characteristics Z, whereas the reservation wage is the same for all: $\overline{w}(X_i) = \overline{w}$. In this way I assume that the decision to participate in the labour market and how much to work is mainly dependent on someone's offered wage.

The expected number of hours worked in the economy can be expressed as follows:

$$\begin{aligned}
Eh &= E(h \mid w \leq \overline{w}) * \Pr(w \leq \overline{w}) + E(h \mid w > \overline{w}) * \Pr(w > \overline{w}) \\
&= 0 * \Pr(w \leq \overline{w}) + E((g(w) + \alpha' Z) \mid w > \overline{w}) * \Pr(w > \overline{w}) \\
&= 0 + \int_{\overline{w}}^{w_{max}} g(w) f(w) dw \frac{\Pr(w > \overline{w})}{\Pr(w > \overline{w})} + E(\alpha' Z \mid w > \overline{w}) * \Pr(w > \overline{w}) \qquad (4) \\
&= \int_{\overline{w}}^{w_{max}} g(w) f(w) dw + E(\alpha' Z \mid w > \overline{w}) * \Pr(w > \overline{w}) \\
&= \int_{w_{min}}^{w_{max}} g(w) f(w) dw - \int_{w_{min}}^{\overline{w}} g(w) f(w) dw + E(\alpha' Z \mid w > \overline{w}) * \Pr(w > \overline{w})
\end{aligned}$$

The second equality follows from my specification of the labour supply in (2) and (3) and the third equality follows from the definition of conditional expectation. For greater clarity I write down the last line of (4) again:

$$Eh = \int_{w_{min}}^{w_{max}} g(w) f(w) dw - \int_{w_{min}}^{\overline{w}} g(w) f(w) dw + E(\alpha' Z \mid w > \overline{w}) * \Pr(w > \overline{w}) \qquad (5)$$

The expression for the average number of hours in the economy (5) allows us to analyse the effects of changes in the inequality in wages on working hours. Since I attempt to analyse changes in working hours purely due to changes in inequality, I want to keep the mean wage unchanged. Therefore, I will analyse the effect of a *mean-preserving spread* in wages on working hours.

Inspecting the expression (5), it turns out that the mean-preserving spread in wages can have an effect on working hours in any direction, either positive or negative. However, in what follows I will argue that the effect of rising inequality on working hours is negative. It is important to remember that I do not claim that this is a general result, or the only possible result. What I do claim, however, is that this result is empirically interesting and relevant. In the empirical section below I will show that the negative effect of wage inequality on working hours prevails in my data.

Let me now turn to the first term on the right hand side of (5); in order to analyse the effect of the mean-preserving spread I refer to Rothschild and Stiglitz (1970 and 1972). They define the mean-preserving spread as a situation where one distribution has been constructed from the other by putting more weight into the tails, while keeping the mean unchanged. This is equivalent to saying that the new variable is distributed as the initial variable plus a white noise. The results from Rothschild and Stiglitz (1970 and 1972) state that for a large class of random variables defined over a bounded interval, as long as *g(.)* is a bounded concave function, $\int_{w_{\min}}^{w_{\max}} g(w)f(w)dw$ will *decrease* when there is a mean-preserving spread in the distribution of w[3].

Intuitively, this stems directly from the concavity of $g(.)$. Because the labour supply is concave, the effect of wages on working hours diminishes with the wage rising. This means that workers with low wages are more responsive to changes in wages than high-paid workers. Now, with the mean-preserving spread in wages, some high-paid workers are getting paid even more, whereas some low-paid workers are getting paid less. Since labour supply is concave, the reduction in hours worked by low-paid workers is large compared to the small increases in hours worked by high earners, hence the average hours in the economy decrease. To further illustrate this point I depict these changes in Figure 1.

Let me turn next to the second term on the right hand side of (5). This term is a residual term that remains after putting the whole wage distribution under the integration in the first term. For wages below the reservation wage the mean-preserving spread pushes the values of $g(.)$ down. That is, offered wages of those who do not participate in the labour market become even smaller. However, at the same time, the number of people out of the labour force increases. Therefore, the resulting effect of this term on working hours is ambiguous. Empirically, it turns out that the negative effect on working hours prevails, as higher variation in wages, keeping the mean wage fixed, decreases working hours in the economy.

Finally, the third term on the right hand side of (5) represents hours worked due to personal characteristics *Z*, conditional on the fact that an individual is working. As the only change in the economy is the mean-preserving spread in wages, there is no change in personal characteristics Z nor is there any change in the reservation wage, $\bar{w}$. Hence, each person that

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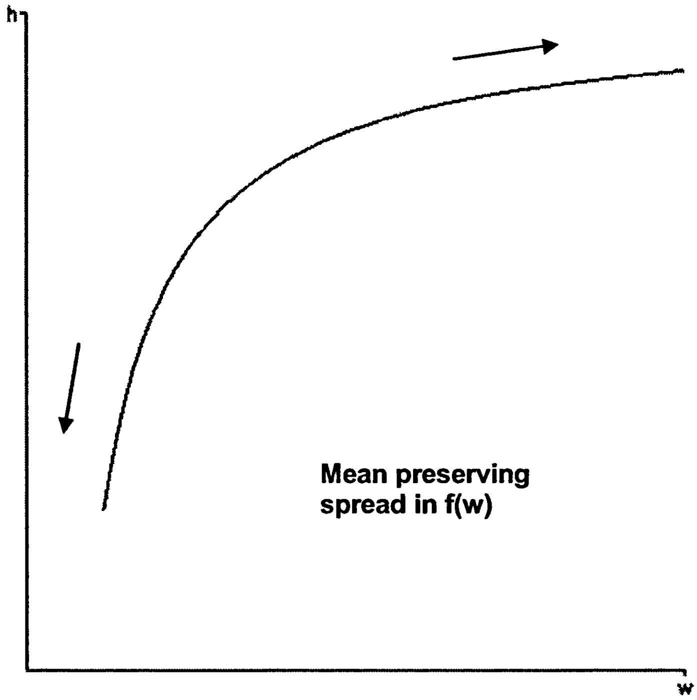
[3] The mean-preserving spread also implies an increase in the variance of *w*. For this reason, in the empirical section below, as a main measure of inequality in wages I will use the standard deviation of log wages. Note, however, that an increase in variance, keeping the mean unchanged, does *not* imply the mean-preserving spread. Nevertheless, I conjecture that for empirically relevant cases of offered wage distributions the implication works in both directions.

remains in the workforce will work exactly the same amount of hours due to Z as before the change in inequality. Perceptive reader would have noticed, however, that the term $E(\alpha' Z \mid w > \overline{w})$ might have changed with the increase in wage inequality. Namely, those who remain in the workforce may on average have a different value of Z from those who opted out. However, in order to ease the discussion I will assume (perhaps unrealistically) that personal characteristics Z are mean independent of w, and thus the only thing that remains to be discussed is the effect on the share of the population that remains active: the probability $\Pr(w > \overline{w})$. If the share of the people in the workforce becomes smaller, this would have as a consequence a negative effect on working hours.

Figure 1: Effect of earnings inequality on working-hours



Theoretically, one cannot say with certainty what happens with the share of the population that participates in the labour force. However, in this chapter I am interested in situations where, after the mean-preserving spread in wages, *more* people opt out of the labour force as their wages fall below the reservation wage. Such situations are, in fact, empirically relevant and have been reported in the literature. Many researchers have found that male participation in the US labour market has been falling over the long term in line with the falling market opportunities. Juhn et al. (1991) report that rising joblessness observed in the 1970s and 1980s is concentrated among groups with declining real wages. Juhn (1992) similarly reports that among prime-age men, declines in employment occur at all ages but are found to be particularly severe among less-educated and low-wage men. Welch (1997) finds that hours worked have fallen most rapidly among the groups that have experienced the greatest reduction in wages. Finally, Juhn and Potter (2006) report that following the decline in demand for less-skilled workers, there has been a decline in wages among less-skilled men and also a sharp fall in the employment of this group. To a large degree, this was characterized with dropping out of the labour force altogether, rather than entering unemployment, which, they argue, has become a predominant response to poor labour market conditions.

For this reason there is one more channel through which hours are reduced in the economy even though higher paid workers increase their labour supply. As assumed above, low-paid workers never supply hours that are very close to zero. In my sample, for example, only about 0.3% of active males work less than 8 hours per week. Hence, when workers decide to leave the labour force, they reduce their working hours to zero from some positive amount and hours are reduced by a discrete interval.

Let me summarize this section. I have shown that much of the empirical labour supply literature supports the idea that labour supply is concave in wage. I will test and confirm this empirically with my data below. The concavity of labour supply in turn implies that increasing wage inequality, holding the average wage constant, decreases working hours in the economy. This can be explained by two effects. First, due to the concavity of the labour supply, working hours' responses are stronger for workers with low and falling wages than for workers with high and rising wages. Second, due to falling wages, some people opt out of the labour force and start supplying zero hours.

3 Data and basic trends

3.1 Data

I use extracts from the CPS Merged Outgoing Rotation Groups (MORG) available from the National Bureau of Economic Research (NBER) web page. The data set covers information on the US labour market over a period of 30 years, from 1979 to 2008, and includes all adult respondents in the outgoing rotation group each month. In my analysis I focus on prime-age men: I restrict the sample to men aged 25-54. By restricting the age band to 25-54 I try to avoid complications due to educational and retirement decisions. I exclude women as female labour supply of the last 30 years has been subject to very strong increases in participation. Furthermore, the labour supply of women is complicated by a multitude of other factors, including child-bearing decisions.

My main variables of interest are hours of work and measures of hourly wages. For the purposes of the analysis, both these variables have undergone certain modifications from the original data set. In what follows I describe this in more details. First I describe the processing of the wage data. Earnings are collected per hour for hourly workers and per week for other workers. Two wage variables are available in the original data set. The first reports the actual hourly wage for hourly workers only. The second reports hourly wage, which comprises of actual hourly wage for hourly workers and imputed hourly wage (weekly earnings divided by usual hours of work) for other workers. Following analogous procedures in Lemieux (2006) and Autor et al. (2008) I force all top-coded values to be equal to \$99 and then multiply all top-coded values by factor of 1.4. I convert wages into 2000 dollars using the quarterly PCE (Personal Consumption Expenditure) Deflator obtained from the US Bureau of Economic Analysis. I put to missing all observations with a wage below \$2.1258 in 2000 dollars (below \$1 in 1979 dollars). I also put to missing all wage observations for self-employed workers. Following Autor et al. (2008) I construct the measure of hourly wage for non-hourly workers by dividing weekly earnings by hours worked last week instead of usual weekly hours. This is done because usual hours are not consistently available over time; there is a break in the series in 1994.

This leaves me with two variables, one measuring hourly wages of workers paid per hour, and one measuring hourly wages for "all" workers. I use these two different measures in the analysis that follows. Wages of hourly workers are the most pure measure of the hourly wage, since they are reported directly by workers themselves. However, the subsample of workers paid per hour is not representative of the whole working population. The measure of hourly wages for "all" workers, on the other hand, is representative of a wider set of workers, but is by construction contaminated by using the information on working hours. If there is a measurement error in the hours worked last week, the constructed hourly wage will be measured with error as well.

These two variables are available only for those workers who are active in the labour market. Therefore, for all those not working (and those whose wage information was put to missing via the process described above) the wage data is missing. Recall that in this chapter I am interested in the behaviour of average hours worked, so I need the information on (offered) wages for both workers and non-workers. To be able to do this and to avoid the sample selection problem, I need to impute wages for all the missing observations. My imputation procedure follows the previous literature of Blau and Kahn (2007), Juhn and Murphy (1997), Juhn (1992) and Juhn et al. (1991), and is done as follows. All the missing observations of wages are imputed from separate regressions by year and by full-time/part-time status. For those not in the labour force, wages are imputed from regressions on the sample of part-time workers[4]. Regressions are done with log wages as the dependent variable and age, age squared, non-white dummy, two education dummies (no high-school as a baseline, high school and college dummy), married dummy and regional and metropolitan status dummies as independent variables. The imputation is done separately for wages of hourly workers and for wages of "all" workers. I thus end up with two basic measures of offered wages in the economy which I call *hourly wage* and *hourly earnings*, respectively.

Now I turn briefly to the measures of labour supply: hours worked per week. My data set contains two variables. The first variable measures usual hours worked and is obtained from the survey question: "How many hours per week does...USUALLY work at this job?", where "this job" stands for the main job of the respondent. Second variable measures hours worked last week, and is obtained from the survey question: "How many hours did...work last week at all jobs?". I set both variables to zero for everybody whose labour status is *not* employed. Originally these observations had hours information missing.

3.2 Basic trends in wages

In this section I show how the main wage variables of interest have evolved over the last thirty years. All the figures and the statistics, such as the mean or the standard deviation of wages, are calculated yearly, from the log real wage variable, and weighted by the earnings weights available from the data source. Recall that these measures include all prime-age men; for those with missing wage data I use imputed wages. Figure 2 shows the average (log) wage of prime-age men over time; as can be seen, hourly wage and hourly earnings move more or less together, only that hourly wage is slightly lower, as hourly workers are on average paid less than the rest of the workers. As has been reported in the literature, the 1980s were a period of falling real wages. Average hourly earnings fell by about 12% in the period from

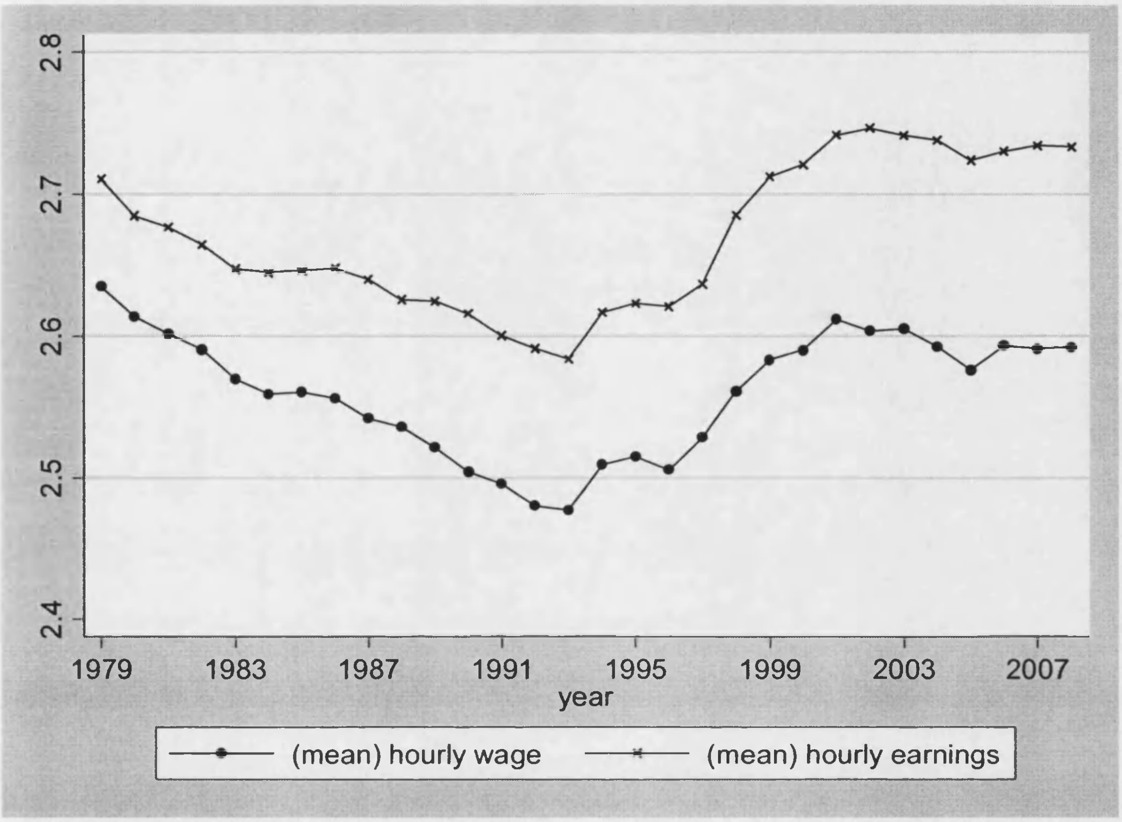
[4] In the literature, wages for non-workers are usually imputed from regressions on workers that work less than a certain number of weeks per year, say 20 weeks. In my data, however, this information is not available. For this reason I use full-time/part-time status as a cut-off.

1979 to about 1993. The 1990s, on the other hand, were a period of rising average wages, with hourly earnings rising by about 17% from 1993 until 2000. Since 2000 real wages have been more or less stable.

Next I show different measures of wage inequality. Figure 3 shows the standard deviation in log real wages over time and Figure 4 shows the ratio of 90^{th} to 10^{th} percentile of log real wage. Consistent with the inequality literature, Lemieux (2006) and Autor et al. (2008), there seems to be quite a steep increase in inequality in the 1980s with flattening of the inequality in the 1990s and after. Note that by the standard deviation measure of inequality, hourly wage and hourly earnings give more or less the same picture for the first 20 years of observations. Since the late 1990s, however, hourly earnings show a slight increase in inequality, whereas the hourly wage measure shows a slight decrease in inequality. Similarly, by the 90/10 percentile ratio measure, the fall of inequality since early 1990s is more pronounced in the hourly wage measure, as compared to the hourly earnings measure.

Finally, I split the wage inequality into the upper-tail and lower-tail inequality. Figure 5 depicts the 90/50 percentile ratio of log real wage, the so-called upper-tail inequality, and Figure 6 depicts the 50/10 percentile ratio of log real wage, the so-called lower-tail inequality. Wage inequality in the upper tail of the wage distribution rose during the 1980s and has more or less flattened or slightly reversed since 1993. Wage inequality in the lower tail rose during the early 1980s and has flattened and reversed since then. According to the two figures, the growth in the 50/10 wage gap has reversed more as compared to the 90/50 wage gap, but the disparity in the evolution of the lower and upper tail inequality does not seem as dramatic as reported in Autor et al. (2008).

Figure 2: Average wage of prime-age men (1979-2008)

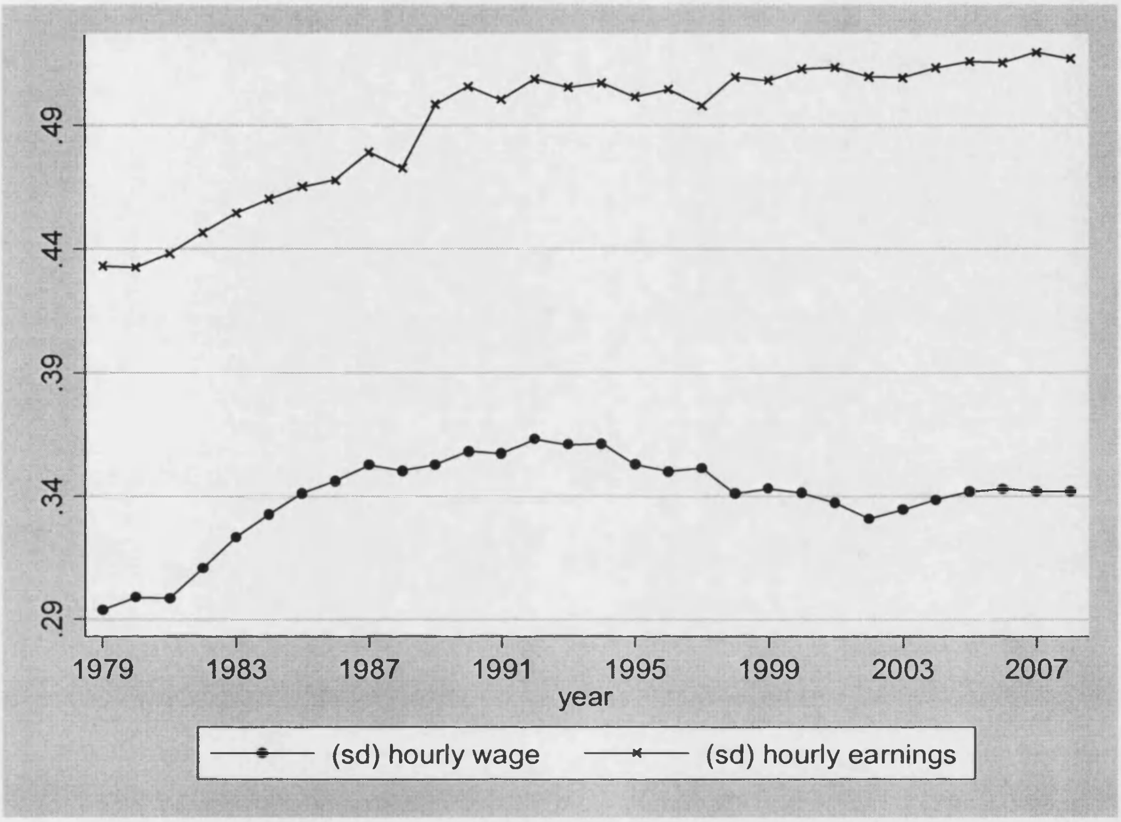


Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. Sample is restricted to men aged 25-54. *Hourly wage*: based on actual hourly wage for hourly workers only. *Hourly earnings*: comprises of actual hourly wage for hourly workers and constructed hourly wage (weekly earnings divided by hours worked last week) for other workers. Plotted wage is the mean log real wage weighted by the earnings weight available in the data source. Units of measurement on the vertical axis are log points.

Following Lemieux (2006) and Autor et al. (2008) I force all top-coded values to be equal to $99 and then multiply all top-coded values by a factor of 1.4. I convert wages into 2000 dollars using the quarterly PCE Deflator obtained from the US Bureau of Economic Analysis. I put to missing all observations with a wage below $2.1258 in 2000 dollars. I also put to missing all wage observations for self-employed workers.

For all those not working and those whose wage information was put to missing via the process described above the wage data is missing. Hence, for these observations I impute the wages. The imputation procedure follows the previous literature, Blau and Kahn (2007), Juhn and Murphy (1997), Juhn (1992) and Juhn et al. (1991) and is done as follows. All the missing observations of wages are imputed from separate regressions by year and by full-time/part-time status. For those not in the labour force, wages are imputed from regressions on the sample of part-time workers. Regressions are done with log wages as the dependent variable and age, age squared, non-white dummy, education dummies, married dummy and regional and metropolitan status dummies as independent variables. Imputation is done separately for hourly wage and for hourly earnings.

Figure 3: Wage inequality: Standard deviation in wages of prime-age men (1979-2008)

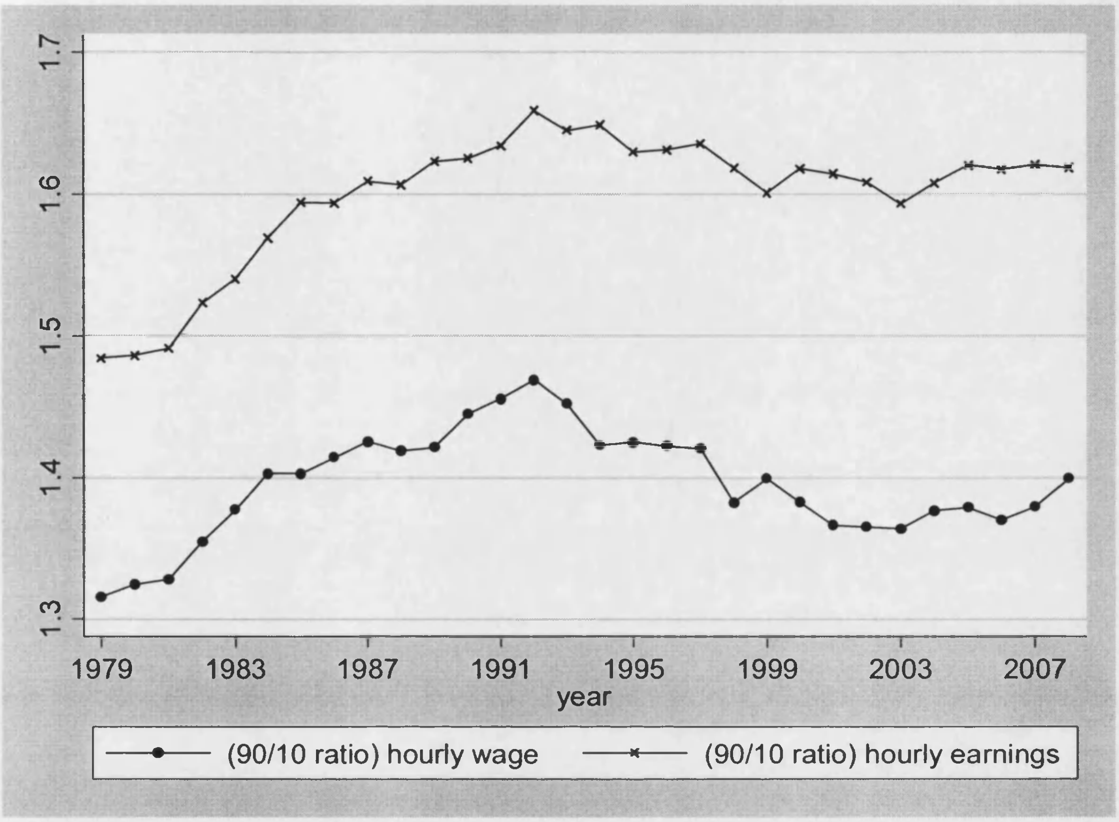


Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. The plotted measure of inequality is the standard deviation of the log real wage. For more details see note under Figure 2. Units of measurement on the vertical axis are log points. The measures of wage inequality show wage inequality in *offered* wages; for individuals that do not work I use imputed wages.

3.3 Wages by quintile in the wage distribution

In this section I show the evolution of wages by wage quintiles. I split workers into 5 different groups according to the quintile of their (offered) wage. Computed average wages from different quintiles are put on the same footing by setting the value of the first observation (year 1979) in each quintile equal to 1. Thus the graph shows the evolution of the average wage in a certain quintile relative to its own value in the year 1979. Similar graphs have been reported in the literature. Juhn et al. (1991), Juhn (1992), Juhn and Murphy (1997), Welch (1997) and Kuhn and Lozano (2008) look at wage changes by deciles or quintiles in the wage distribution. Similarly, Blundell and MaCurdy (1999), Welch (1997) and Autor et al. (2008) report the evolution of wages by education groups. My findings are in line with the literature.

Figure 4: Wage inequality: 90/10 percentile ratio in wages of prime-age men (1979-2008)



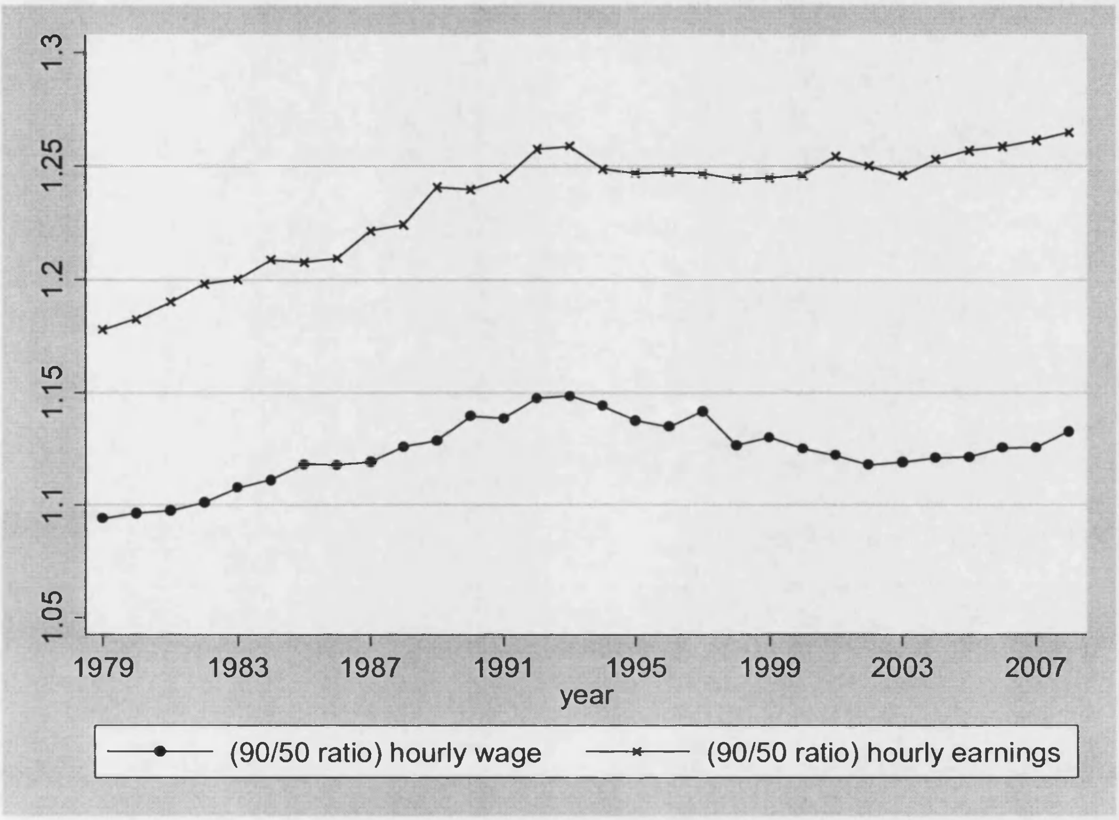
Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. The plotted measure of inequality is the ratio of 90th and 10th percentile of the log real wage. For more details see note under Figure 2. Units of measurement on the vertical axis are log points. The measures of wage inequality show wage inequality in *offered* wages; for individuals that do not work I use imputed wages.

In Figure 7 I show the evolution of the hourly wage and in Figure 8 I show the evolution of the hourly earnings by wage quintiles. The two pictures tell more or less the same story, which is also consistent with the story from the graphs on overall wage inequality. In the 1980s, whereas real (offered) wages of the top two wage quintiles were basically stable, wages of workers with wages lower than the 60th percentile fell. The steepest falls were experienced by the lowest paid workers in the first and second quintile. Looking at the hourly earnings in Figure 8, real wages for workers in the first and second wage quintile fell by about 19% from 1979 to 1992![5] This of course implies that inequality was on the rise in the 1980s at the expense of low-paid workers. One can clearly notice the divergence of the three lines towards the middle of the picture, indicating an increase in inequality. At the beginning of the 1990s this trend started to reverse. The five lines started to get closer together, as the wages of

[5] Note that this cannot be read directly form the graph, because units of measurement are in terms of log ratios. I calculated this directly from the data.

the lowest paid workers started to catch up with the general increase in average wages. Depending on whether one looks at the hourly wage or hourly earnings picture, wage inequality seems to have slightly decreased or flattened by the end of the period, respectively.

Figure 5: Wage inequality: 90/50 percentile ratio in wages of prime-age men (1979-2008)



Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. The plotted measure of inequality is the ratio of 90^{th} and 50^{th} percentile of the log real wage. For more details see note under Figure 2. Units of measurement on the vertical axis are log points. The measures of wage inequality show wage inequality in *offered* wages; for individuals that do not work I use imputed wages.

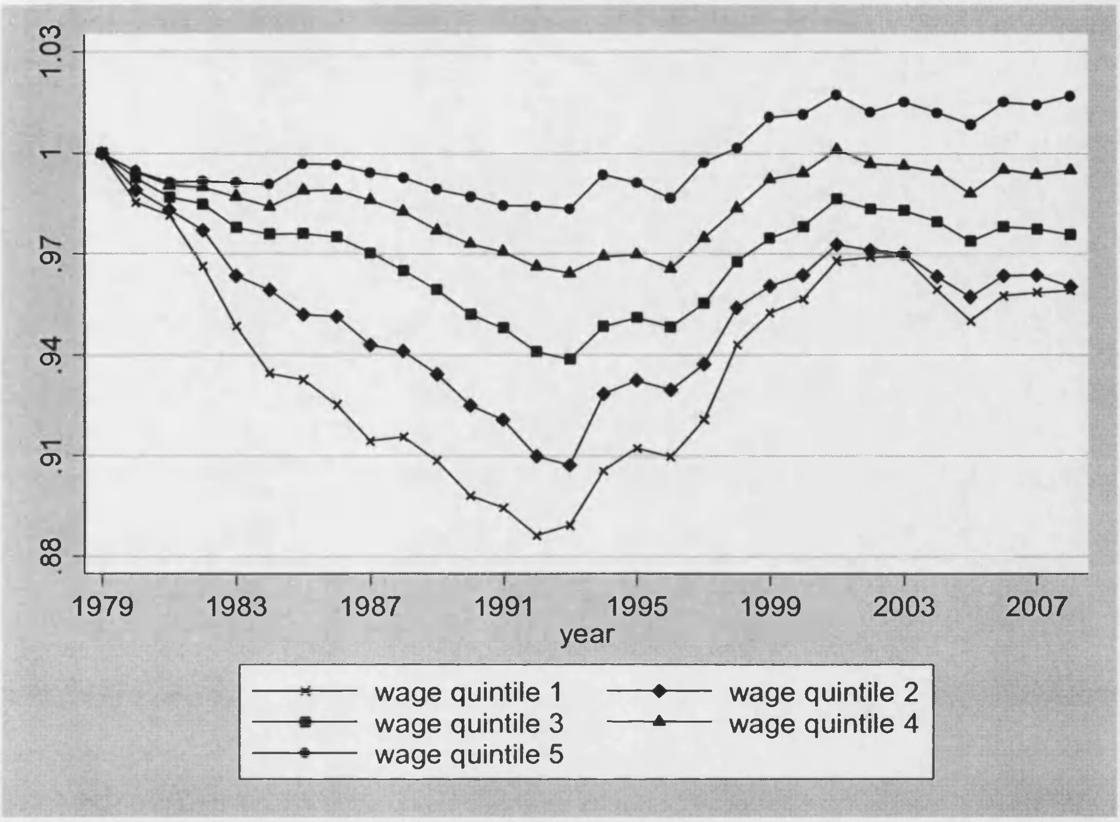
This graph is suggestive of a simple story that I am trying to tell in this chapter. Take for example the inequality increase from the 1980s. We can clearly see how the wages of low-paid workers were falling whereas wages of high-paid workers remained largely unchanged. Now what does this imply for the labour supply? It implies, as already suggested by Juhn et al. (1991), Juhn (1992), Welch (1997) and Juhn and Potter (2006), that the labour supply of high-paid workers remained stable, while the labour supply of low-paid workers must have fallen significantly, since their market opportunities diminished. Similar logic can be used for the later period, when the wages of all groups were rising. Now, all groups should be putting more effort into the labour market as they are being rewarded with higher wages.

Figure 6: Wage inequality: 50/10 percentile ratio in wages of prime-age men (1979-2008)

Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. The plotted measure of inequality is the ratio of 50th and 10th percentile of the log real wage. For more details see note under Figure 2. Units of measurement on the vertical axis are log points. The measures of wage inequality show wage inequality in *offered* wages; for individuals that do not work I use imputed wages.

So far, however, there has been no discussion of wage inequality per se, and all these labour supply effects could in aggregate simply be attributed to the rising and falling of the *average* wage in the economy. The question therefore is what happens if one keeps the average wage constant. Suppose again that the spreading out of wages by wage quintiles would be as it was in the 1980s, but the average wage in the economy would remain constant – a mean-preserving spread in wages. That is, wages would fall for low-paid workers and rise for high-paid workers. In this case I argue that due to the concavity of the labour supply, working hours of low-paid workers would decrease more than the working hours of high-paid workers, and average working hours in the economy would fall. In the next section I turn to the analysis of working hours and their dependence on wages and wage inequality.

Figure 7: Average hourly wage by wage quintiles of prime-age men (1979-2008)



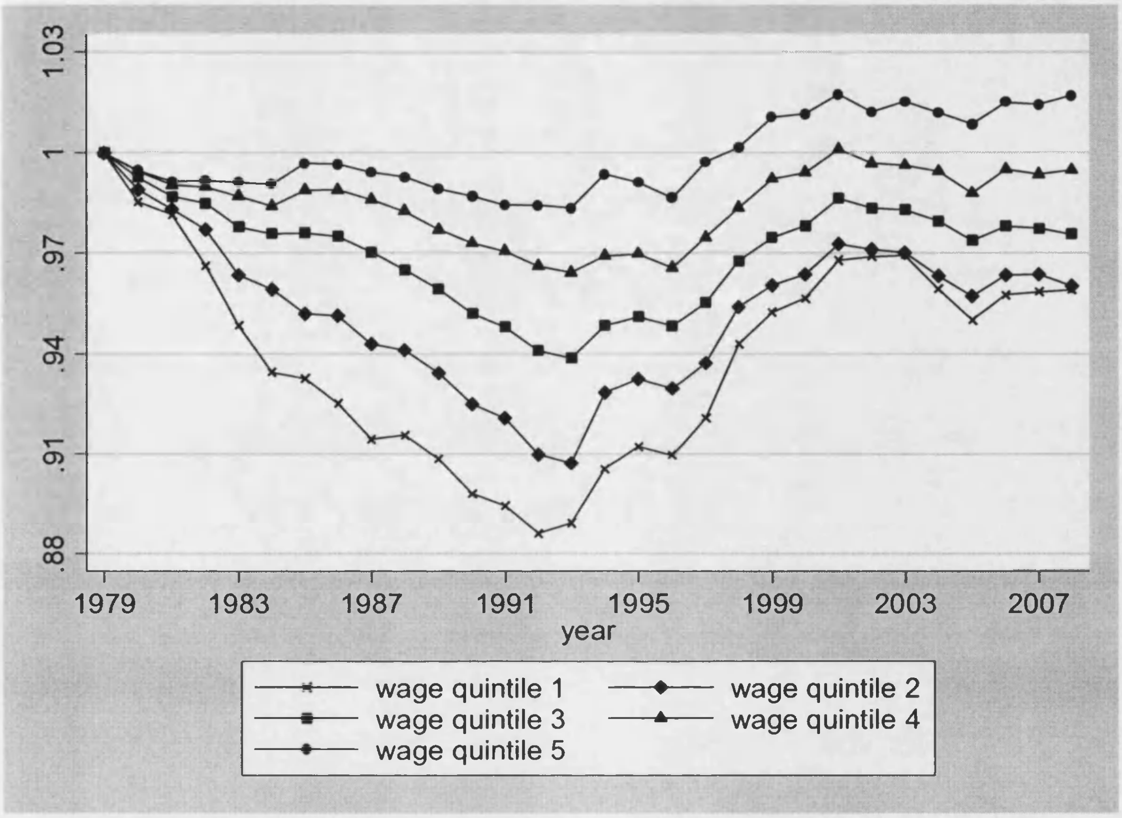
Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. *Hourly wage*: based on actual hourly wage for hourly workers only. The graph shows the evolution of average log real wage by quintiles of the wage distribution. "wage quintile 1" represents the first (lowest) quintile and so on, with the "wage quintile 5" representing the highest fifth quintile. Computed average wages from different quintiles are put on the same footing by setting the value of the first observation (year 1979) equal to 1. Thus the graph shows evolution of the log average wage in a certain quintile relative to the value of the log real wage in that quintile in the year 1979. Units of measurements on vertical axis are in terms of log ratios! For more details about the wage measure see note under Figure 2. The measures are based on *offered* wages; for individuals that do not work I use imputed wages.

4 Results

4.1 Working hours and the concavity of labour supply

In Figure 9 I show average weekly hours of work over time. It is obvious from Figure 9 that there is a strong cyclical component in working hours over time, closely following the (opposite) movements in the aggregate unemployment rate (not shown). I could regress out the cyclical component, or try to derive a trend measure of working hours. However, working hours depend on several things, such as the economic cycle, the unemployment rate, the age composition of the population etc. The choice of controlling for any of these for the purposes of graphical representation would be arbitrary, so I decided to show a crude measure of working hours over time instead. In the regression analysis below, I will of course control for the unemployment rate and age.

Figure 8: Average hourly earnings by wage quintiles of prime-age men (1979-2008)



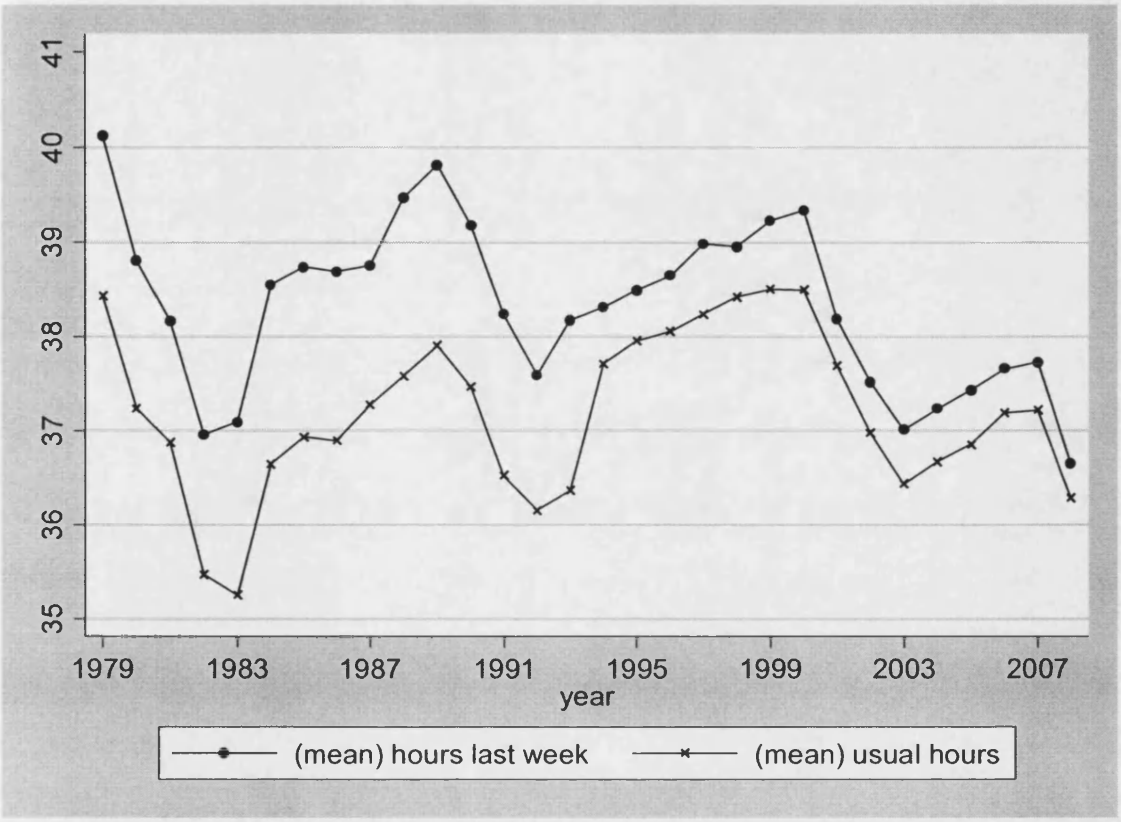
Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. *Hourly earnings*: comprises of actual hourly wage for hourly workers and constructed hourly wage (weekly earnings divided by hours worked last week) for other workers. Units of measurements on vertical axis are in terms of log ratios! For more details see note under Figure 7. The wage measures are based on the distribution of *offered* wages; for individuals that do not work I use imputed wages.

In Figure 10 I show the evolution of working hours over time by quintiles of hourly earnings. The measure of working hours used here is hours worked last week. Due to an apparent break in the series in 1994, I regress the hours' measures in each wage quintile on a dummy variable whose value is set to 0 before 1994 and to 1 in 1994 and after. Residuals from these regressions are put on the same footing by setting the starting value to zero[6]. One can clearly see from the picture how working hours of low-paid workers are more responsive than of the rest. However, from a graphical analysis it is not possible to say anything about the response of the labour supply of different quintile groups to wages. Most likely, the obvious patterns in

[6] Autor et al. (2008) mention the break in the series for the usual hours worked in the year 1994. However, after inspecting graphs of working hours over time by quintiles, it becomes apparent that there is some kind of break also in the hours worked last week variable. It is not clear to me where this break comes from.

the picture follow the evolution of the economic cycle – showing large responsiveness of low-paid workers' employment to economic shocks.

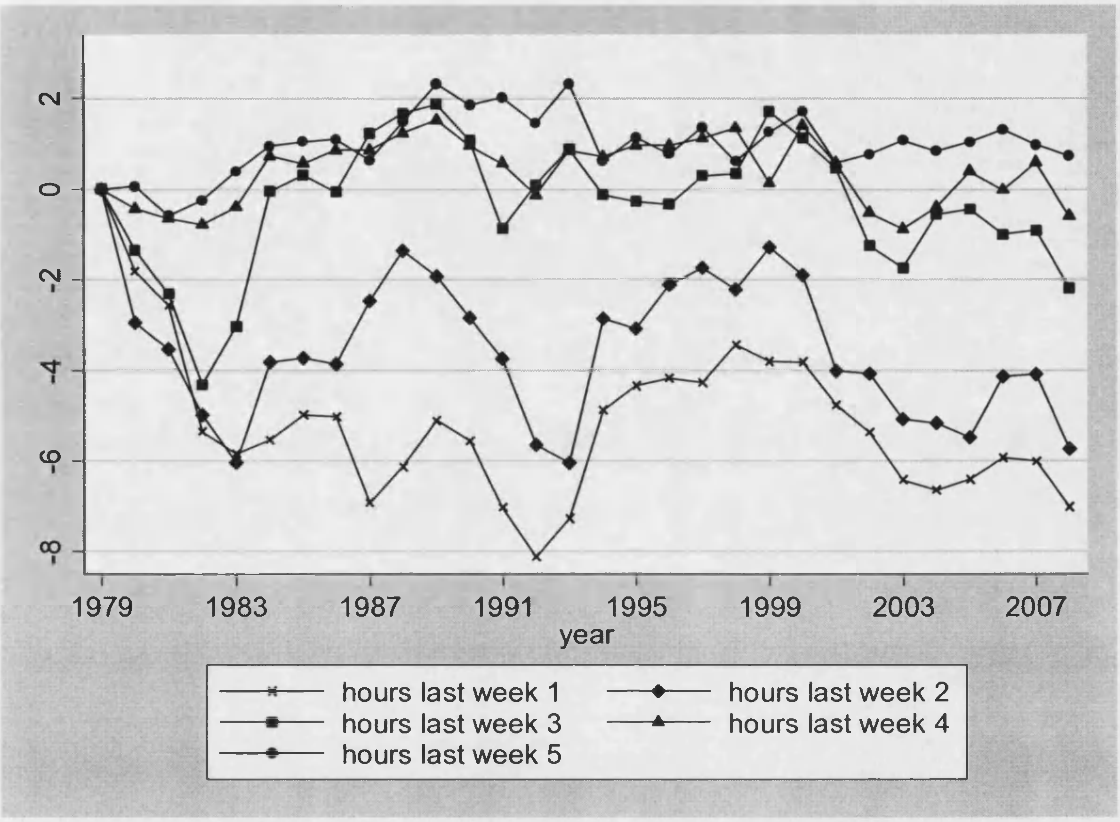
Figure 9: Average hours of work per week of prime-age men over time (1979-2008)



Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER web page. *Hours last week* measure is obtained from the survey question: "How many hours did...work last week at all jobs?". *Usual hours* measure is obtained from the survey question: "How many hours per week does...USUALLY work at this job?", where "this job" stands for the main job of the respondent. I set both variables to zero for everybody whose labour status is not employed; originally these observations had hours information missing. Graph is for men aged 25-54. There is an apparent break in the series in 1994.

I next turn to testing whether labour supply is indeed concave in wages. In order to test this hypothesis I regress individual working hours on wage and wage squared. If labour supply is concave, I would expect the coefficient on the wage level to be positive and the coefficient on the wage squared to be negative. In all regressions I include year dummies, age, age squared, a non-white dummy, a married dummy, regional dummies, a metropolitan status dummy and a constant. I am in effect estimating a cross-sectional labour supply function. Note that education is excluded from my empirical model; following Juhn et al. (1991) I assume that more educated persons work more because they earn higher wages. Hence, the main measure of marketable skill in the labour market is someone's (offered) wage.

Figure 10: Average hours worked last week by wage quintile of prime-age men (1979-2008)



Source: own calculations based on CPS Merged Outgoing Rotation Groups obtained from the NBER webpage. This picture shows evolution of working hours over time by quintiles of hourly earnings. The measure of working hours used here is *hours last week*. Due to apparent break in the series in 1994, I regress hours' measures in each wage quintile on a dummy variable, whose value is set to 0 before 1994 and to one 1 in 1994 and after. Residuals from these regressions are put on the same footing by setting the starting value to zero. For more details see note under Figure 9.

Results are reported in Table 1. I show four different specifications, with two different measures of working hours (hours last week and usual hours worked) and two different measures of wages (hourly wage and hourly earnings). Due to potential endogeneity of the wage variable, I report results from both OLS and Instrumental variables (two stage least squares) regressions. To instrument for wage and wage squared I use wage decile dummies as instruments, indicating in which decile of the wage distribution an individual's wage offer falls (Baker & Benjamin (1997), Juhn & Murphy (1997), Blau et al. (2003), Blau & Kahn (2007)). Using wage decile dummies can potentially correct for measurement error in wage.

Table 1: Concavity of labour supply for prime-age men

Dependent variable:	hours worked last week		usual hours worked	
	(1)	(2)	(3)	(4)
OLS				
hourly wage	1.561		1.396	
	(0.004)***		(0.004)***	
hourly wage squared	-0.013		-0.011	
	(0.000)***		(0.000)***	
hourly earnings		0.528		0.544
		(0.003)***		(0.003)***
hourly earnings squared		-0.006		-0.004
		(0.000)***		(0.000)***
Instrumental variables				
hourly wage	3.900		3.461	
	(0.014)***		(0.013)***	
hourly wage squared	-0.078		-0.068	
	(0.000)***		(0.000)***	
hourly earnings		0.871		0.712
		(0.006)***		(0.005)***
hourly earnings squared		-0.013		-0.007
		(0.000)***		(0.000)***

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include year dummies, regional dummies, metropolitan status dummy, age, age squared, non-white dummy, married dummy and a constant. In instrumental variables regressions wage and wage squared are instrumented using wage decile dummies.

In all specifications there is evidence of a positive and concave effect of wages on working hours. The effect of the wage level is positive and the effect of wage squared is negative, as predicted. All coefficients are very precisely estimated. In confirmation with the attenuation bias hypothesis, wage effects are larger in absolute value in the instrumental variables specification. To give a feel for the extent of concavity, I compute wage effects for the specification reported in the column (2) in Table 1. Based on the OLS (IV) regression results, compared to the wage effect at the 10^{th} percentile, the effect of wage at the median (50^{th} percentile) is 8 (12) % smaller, while the effect at the 90^{th} percentile is 23 (32) % smaller. Therefore, I conclude that labour supply is concave in wage and that at the top of the wage distribution the effects of wages are about one third smaller than at the bottom of the wage distribution.

In the preceding analysis I basically ignored the fact that for many observations the value of the dependent variable is equal to zero. That is, some prime-age men do not work. In my sample, about 12% of men aged 15-54 report zero hours worked. In order to check whether my results are robust to alternative specifications in which this fact is accounted for explicitly I do the analysis using two alternative specifications[7]. First, in order to take into account the

[7] Results are not reported here but can be obtained upon request.

mass of observations at zero hours, I use a Tobit instead of the linear regression. However, as expected with male labour supply, the Tobit specification gives very similar results. Comparing the results of the Tobit specification with results in the top panel of Table 1 confirms the results reported above. Second, following alternative specifications in Blau and Kahn (2005, 2007), I assign missing wages using Heckman (1979) selectivity bias correction. I estimate the wage equation following the Heckman maximum likelihood estimation with age and age squared, a non-white dummy, two dummies for education, metropolitan area, year and region dummies on the right hand side, and a married dummy included as an extra identification variable in the selection equation. The resulting wage equation is then used to predict wages for all those whose wage data was missing. With this new data for wages I then do the same analysis as above and the results are again similar.

4.2 Effects of wage inequality on working hours

In this section I explore the effects of wage inequality on the labour supply in the economy. I argued theoretically in section 2.2 that average hours of work become lower with a mean-preserving spread in wages. In other words, a rise in wage inequality, holding the average wage constant, decreases average hours of work. Since the basic theoretical idea refers to averages and aggregate measures of variation, I first test it using aggregated data. This is obtained by collapsing the micro-data set by years, computing means and measures of inequality using earnings weights. By its nature, this is a time-series data set and identification is obtained by exploiting the variation over time.

The results are reported in Table 2. Again, I report four different specifications, with two different measures of working hours (hours last week and usual hours worked) and two different measures of wages (hourly wage and hourly earnings). The dependent variable (average working hours in the economy) is regressed on average log wage and standard deviation of log wages. Furthermore, I include the aggregate unemployment rate and average age. The first is used to control for the involuntary unemployment and the business cycle while the second is used to control for effects of the age composition of the population on working hours. In the last thirty years there has been significant aging of the (prime-age) male workforce, which may have effects on its own. Note that due to a break in the hours' series in 1994, I first use an auxiliary regression of working hours on a dummy variable, whose value is set to 0 before 1994 and to 1 in 1994 and after. In the main regressions I use the residuals from the auxiliary regression on the left hand side as measures of working hours.

Table 2: Effects of wage inequality on working hours – aggregate regressions

Dependent variable:	(mean) hours last week		(mean) usual hours	
	(1)	(2)	(3)	(4)
OLS				
(mean) hourly wage	-6.190		-7.434	
	(3.090)*		(2.385)***	
(sd) hourly wage	-23.547		-28.033	
	(7.207)***		(5.563)***	
(mean) hourly earnings		-5.425		-5.332
		(2.393)**		(2.029)**
(sd) hourly earnings		-19.201		-20.317
		(4.908)***		(4.160)***
unemployment rate	-0.833	-0.820	-0.794	-0.770
	(0.058)***	(0.055)***	(0.044)***	(0.046)***
(mean) age	-0.826	-0.322	-0.738	-0.235
	(0.098)***	(0.199)	(0.076)***	(0.169)
Breusch-Godfrey AR(1) test (p-value)	0.01	0.01	0.01	0.05
MLE AR(1)				
(mean) hourly wage	-1.362		1.173	
	(3.817)		(2.565)	
(sd) hourly wage	-13.329		-10.941	
	(7.179)*		(5.860)*	
(mean) hourly earnings		-1.342		3.700
		(3.921)		(2.639)
(sd) hourly earnings		-11.961		0.516
		(5.337)**		(4.048)
unemployment rate	-0.780	-0.782	-0.715	-0.694
	(0.057)***	(0.051)***	(0.047)***	(0.056)***
(mean) age	-0.904	-0.637	-0.829	-0.623
	(0.176)***	(0.269)**	(0.172)***	(0.424)
Observations	30	30	30	30

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%; Constant included in all regression but not reported. Breusch-Godfrey test for AR(1) process in disturbances; p-values based on F-statistic. The wage measures are based on the distribution of *offered* wages; for individuals that do not work I use imputed wages.

In the top panel of Table 2 I report simple OLS results. In general, the results support the idea that higher inequality has a negative effect on working hours, after controlling for average wage, unemployment rate and average age. The coefficient is negative and highly significant in all four specifications. This is a good result. Furthermore, the coefficients in front of unemployment rate and average age have the expected negative and significant signs. What is puzzling, however is the negative (and significant) sign of the coefficient in front of the average wage variable. According to theory, this coefficient should be positive. It is possible however, that there is a time component that my model does not capture. Therefore, I test for the AR(1) process in the errors by using Breusch-Godfrey test. In all four specifications, the null hypothesis of *no* AR(1) process is rejected at high significance. Therefore, in the bottom panel of Table 2 I report results of the Maximum Likelihood estimation with the AR(1) process in the error term. The negative effect of wage inequality on working hours is again confirmed in three out of four cases (at lower significance, though), and the coefficient on average wage becomes insignificant.

Now I turn to the analysis of the individual data. I test the same idea as before: What is the effect on working hours of individuals of the change in the standard deviation of wages, keeping the average wage constant. Hence I regress individual working hours on the *aggregate* measure of average log wage and the measure of wage inequality. Individual data allow me to also control for some personal characteristics. Results are reported in Table 3 and Table 4. In Table 3 I use standard deviation in log wages as a measure of wage inequality, while in Table 4 I use the 90/10 percentile ratio of log wages. Other controls include aggregate unemployment rate and individual controls such as age, age squared, non-white dummy and married dummy. All regressions also include regional and metropolitan status dummies. Because regressions are done on individual data but using aggregate regressors (average log wage, log wage inequality and unemployment rate) I report clustered standard errors, clustered by year.

Table 3: Effects of wage inequality on working hours (I)

Dependent variable:	hours worked last week		usual hours worked	
	(1)	(2)	(3)	(4)
(mean) hourly wage	-15.490		-8.439	
	(3.601)***		(1.690)***	
(sd) hourly wage	-32.989		-22.633	
	(9.709)***		(3.889)***	
(mean) hourly earnings		-7.353		-2.516
		(0.846)***		(1.011)**
(sd) hourly earnings		-16.490		-7.761
		(1.859)***		(1.968)***
unemployment rate	-0.637	-0.769	-0.770	-0.784
	(0.074)***	(0.044)***	(0.039)***	(0.051)***
age	1.053	1.061	0.948	0.950
	(0.034)***	(0.034)***	(0.024)***	(0.024)***
age squared	-0.015	-0.015	-0.013	-0.013
	(0.000)***	(0.000)***	(0.000)***	(0.000)***
non-white dummy	-5.363	-5.342	-5.113	-5.109
	(0.190)***	(0.189)***	(0.116)***	(0.116)***
married dummy	6.849	6.811	6.670	6.662
	(0.055)***	(0.055)***	(0.077)***	(0.076)***
Constant	71.607	49.112	52.283	33.549
	(12.513)***	(2.816)***	(5.422)***	(3.168)***
Observations	2318780	2318780	2148705	2148705
R-squared	0.05	0.05	0.06	0.06

Robust standard errors in parentheses (clustered by year), * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include regional dummies and metropolitan status dummy. The wage measures are based on the distribution of *offered* wages; for individuals that do not work I use imputed wages.

There is very strong evidence of a negative effect of wage inequality on working hours. Effects in all specifications except one are negative and highly significant. While this result is reaffirming, the puzzling negative and significant effect of average log wage on working hours remains. One potential reason for this is because non-earned income is not included as a control into my regressions. If non-earned income is positively correlated with average wages in the economy then this could explain the negative coefficient. However, this information is not available in my data. For this reason I will do an analogous analysis with an alternative data set, where this information is available. See section 4.3 below.

Table 4: Effects of wage inequality on working hours (II)

Dependent variable:	hours worked last week		usual hours worked	
	(1)	(2)	(3)	(4)
(mean) hourly wage	-11.682		-9.815	
	(5.558)**		(2.864)***	
(90/10 ratio) hourly wage	-8.519		-10.849	
	(6.709)		(2.986)***	
(mean) hourly earnings		-10.921		-4.375
		(1.126)***		(1.137)***
(90/10 ratio) hourly earnings		-9.314		-4.817
		(1.473)***		(0.951)***
unemployment rate	-0.489	-0.756	-0.685	-0.787
	(0.055)***	(0.045)***	(0.036)***	(0.045)***
age	1.049	1.059	0.945	0.950
	(0.033)***	(0.034)***	(0.023)***	(0.024)***
age squared	-0.014	-0.015	-0.013	-0.013
	(0.000)***	(0.000)***	(0.000)***	(0.000)***
non-white dummy	-5.373	-5.345	-5.117	-5.109
	(0.189)***	(0.190)***	(0.115)***	(0.116)***
married dummy	6.871	6.816	6.680	6.663
	(0.056)***	(0.055)***	(0.078)***	(0.076)***
Constant	61.715	65.404	62.787	42.435
	(23.237)**	(5.118)***	(11.274)***	(4.025)***
Observations	2318780	2318780	2148705	2148705
R-squared	0.05	0.05	0.06	0.06

Robust standard errors in parentheses (clustered by year), * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include regional dummies and metropolitan status dummy. The wage measures are based on the distribution of *offered* wages; for individuals that do not work I use imputed wages.

To check the robustness of the results I do the analysis with the two alternative specifications mentioned above. The Tobit specification gives results that are in line with the conclusions drawn above. Comparing the results of the Tobit specification with the results in Table 3 and Table 4 shows that the results are very similar. Second, I use wages that were imputed using the Heckman procedure. When I compare the results of the analysis on aggregated data (analogous to Table 3), in two out of eight cases the results are statistically significant with a negative sign, in six other cases the results are not statistically significant, but in five cases the sign is negative. Turning to individual data (analogous to Table 3 and Table 4), in six out of eight cases the results are similar. In the other two cases, the results are not statistically significant but exhibit the correct sign. To summarize, the two alternative specifications, a Tobit analysis and a Heckman procedure for missing wages, give the same conclusions as the main analysis in most cases and support the hypothesis that increased wage inequality reduces working hours.

Let me now give some perspective on the empirical importance of the effect of inequality on working hours. I compute the magnitude of the effects of the change in wage inequality on working hours. For example, the results from column (2) in Table 3 with hourly earnings as a measure of wages and working hours last week on the left hand side report a value of -16.490 for the coefficient on the standard deviation of log wage. What is the significance of this in practice? From 1979 to 2008, the standard deviation in log hourly earnings rose from 0.43 to about 0.52. This would cause, assuming that the average wage[8] and other things did not change over time, a 1.4 hours decrease in hours worked last week, which represents a 3.6% decrease in working hours from the year 1979. Similarly, doing the same computation for the specification in column (1) of Table 3, the increase in the standard deviation of log hourly wage would represent a 1.6 hours (4%) decrease in hours worked last week. The magnitudes are slightly lower if I compute the effect of the standard deviation in wages on usual hours worked (columns (3) and (4) of Table 3). For the same time period, standard deviation in log hourly wages would decrease usual hours worked by 1.1 hours (2.8%), whereas the standard deviation in log hourly earnings would decrease usual hours worked by 0.7 hours (1.7%). To give more perspective, note that in the US over the period from 1979 to 2008, hours worked last week decreased by about 8.6% and usual hours decreased by about 5.6% on average. From this it follows that, if we take the results at face value, the rising inequality could potentially explain 30-40% of the decrease in working hours over time.

The empirical analysis so far has been supportive of the idea that greater wage inequality, keeping the average wage constant, decreases working hours. But the analysis has been done with the measures of wages *gross* of taxes. One could argue that what workers really care about in the labour market is their wage *net* of taxes. It is important to note that there have been considerable changes in marginal tax rates in the US during the period analysed in this chapter, and hence the results may be biased. However, I rely on findings reported in Mroz

[8] From 1979 to 2008 the average real wage did in fact change very little, if one looks only at the beginning period and the end period.

(1987), Devereux (2004) and Blau and Kahn (2007). Mroz estimates labour supply of married women and compares results from the model without taxes to the model with taxes. He concludes that the influence of taxes on the estimates of the labour supply parameters appears to be at most of a second order importance. Similarly, Devereux (2004) and Blau and Kahn (2007) in their robustness checks, use some simplifying assumptions to compute after-tax wages from their pre-tax wages data, and find that an analysis with after-tax data yields very similar results. I conclude that using wages gross of taxes does not bias my results very much.

4.3 Effect of wage inequality - IPUMS-CPS data

In the preceding section I find evidence that wage inequality, ceteris paribus, has a negative effect on working hours in the economy. However, this came along with a puzzling result that the coefficient in front of the average wage variable was negative. For this reason I do analogous analysis on alternative data obtained from the IPUMS CPS data base (King et al. (2004)). The time span of this data is shorter, going back only to the year 1990, but there are some added advantages: there is information on the non-earned income of individuals as well as information on the number of *weeks* worked in a (previous) year.

I use the same sample restrictions and processing of the data set as described in section 3.1 above, the only exception being the way in which I impute wages for individuals not in the labour force. Recall that in the literature (Blau and Kahn (2007), Juhn and Murphy (1997), Juhn (1992) and Juhn et al. (1991)) wages for non-workers are usually imputed using the wage regressions for people who work less than a certain number of weeks per year. Previously, this information was not available; however, now that it is available, I impute wages for non-workers based on the estimated wage equation for workers that worked less than 20 weeks in the previous year.

Table 5 reports results from regressions on individual data, with average log hourly earnings as a measure of average wage and standard deviation of log hourly earnings as a measure of wage inequality[9]. I report four different specifications, depending on what variable is used on the left hand side to measure working hours: hours worked last week, usual weekly hours worked last year, share of weeks worked last year and the product of the last two (usual weekly hours times the share of weeks worked last year). Other controls include the aggregate unemployment rate and individual controls such as age, age squared, non-white dummy and

[9] I only report results for hourly earnings because results using hourly wage as a measure of wages produces qualitatively similar results.

married dummy. All regressions include regional dummies. One additional control that was not available in the CPS MORG data is the measure of non-earned income of the individual.

Results in Table 5 support the theoretical prediction that wage inequality decreases hours worked in the economy. In all four specifications the coefficient in front of the wage inequality variable is negative and highly significant. Note that in this case the coefficient in front of the wage level variable is not significantly different from zero, which is reassuring. Other coefficients have the expected signs. The analysis on this alternative data set therefore provides one more piece of evidence for the idea that the mean-preserving spread in wages lowers hours worked in the economy.

Table 5: Effects of wage inequality on working hours (IPUMS-CPS data set)

Dependent variable:	hours worked last week	usual weekly hours worked (last year)	share of weeks worked (last year)	usual weekly hours * share of weeks worked (last year)
	(1)	(2)	(3)	(4)
(mean) hourly earnings	-1.964	0.638	0.074	3.627
	(2.271)	(2.433)	(0.055)	(2.725)
(sd) hourly earnings	-32.168	-35.615	-0.319	-26.609
	(7.608)***	(7.082)***	(0.163)*	(7.365)***
unemployment rate	-0.707	-0.209	-0.009	-0.491
	(0.106)***	(0.092)**	(0.002)***	(0.096)***
non-earned income (000)	-0.270	-0.236	-0.007	-0.282
	(0.020)***	(0.017)***	(0.000)***	(0.021)***
age	0.918	0.703	0.010	0.903
	(0.078)***	(0.066)***	(0.001)***	(0.070)***
age squared	-0.012	-0.010	-0.000	-0.012
	(0.001)***	(0.001)***	(0.000)***	(0.001)***
non-white dummy	-4.834	-4.203	-0.078	-4.668
	(0.233)***	(0.194)***	(0.003)***	(0.204)***
married dummy	7.080	5.766	0.116	6.668
	(0.129)***	(0.148)***	(0.003)***	(0.152)***
Constant	43.062	43.021	0.663	24.558
	(5.926)***	(6.651)***	(0.170)***	(8.243)***
Observations	140060	140060	140060	140060
R-squared	0.06	0.07	0.09	0.09

Robust standard errors in parentheses (clustered by year), * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include regional dummies. The wage measures are based on the distribution of *offered* wages; for individuals that do not work I use imputed wages.

To check the robustness of the results, I do the analysis also using Tobit specification. In this data about 15% of men report zero hours worked, and as expected, results with Tobit specification are very similar to results reported in Table 5. When checking the results using Heckman selectivity bias correction, I estimate the wage equation with age and age squared, a non-white dummy, two dummies for education, metropolitan area dummies, year dummies, and region dummies on the right hand side. The selection equation includes non-earned income and a married dummy as additional regressors. The resulting wage equation is then used to predict wages for all those whose wage data was missing. With the new data on wages, results are again similar to those reported in Table 5, providing additional evidence that a mean-preserving spread in wages has a negative effect on the hours worked in the economy.

5 Conclusion

In this chapter I argue that a mean-preserving spread in offered wages (rising wage inequality) lowers average working hours in the economy. The exact mechanism works through the concavity of the labour supply. Intuitively, with wage inequality rising, the wages of highly paid workers rise and wages of low-paid workers fall. Due to the concavity in the labour supply, responses to the changes in wages are stronger at the bottom of the wage distribution, hence working hours of low-paid workers decrease more than working hours of high-paid workers. Some of the low-paid workers may also decide to drop out of the labour force and become inactive, thus reducing their labour supply to zero. Therefore, as a consequence, average working hours in the economy fall.

This explanation implicitly assumes that working hours are just a response to labour market opportunities. In this I follow the labour supply literature such as Juhn et al. (1991), Juhn (1992) and Welch (1997). The contribution of this chapter is that it explicitly recognizes that changes in the inequality of offered wages have effects on average working hours in the economy. Using the CPS-MORG data for prime-age men I find evidence for this explanation. First I establish empirically the concavity of the labour supply function. Secondly, using various variables to measure wages and labour supply, and using several different specifications, I find empirical evidence that after controlling for the average wage, wage inequality has indeed a negative and significant effect on the labour supply. According to my results, the increase in wage inequality that occurred in the US over the 1979-2008 period would cause a 3% decrease in average hours worked in the economy. Rising inequality can thus potentially explain 30-40% of the decrease in working hours over time.

The results in this chapter are interesting for students of labour supply, for students of inequality and for researchers interested in explaining differences in working hours across countries and over time. Moreover, the results are also interesting from the point of view of public policies that affect wage inequality.

6 References

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Chapter III

Do Windfall Gains Affect Labour Supply? Evidence from the European Household Panel

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Abstract: We investigate whether workers adjust hours worked in response to windfall gains using data from the European Household Panel from 1994 to 2001. The results suggest that unexpected variation in income has a negative (although small) effect on working hours. In particular, after receiving an unanticipated windfall gain, individuals are more likely to drop out of the labour force and the effects become larger as the size of windfall increases. Furthermore, the empirical findings show that the impact of windfall gains on labour supply: (i) is more important for young and old individuals, (ii) is most negative for married individuals with young children, (iii) but can be positive for single individuals at the age of around 40 years.

1 Introduction

What is the effect of windfall gains on economic behaviour? A popular belief presumes that the majority of people would quit work if they won a lottery. But do windfall gains have an impact on individuals' working hours? According to the life-cycle model, a relaxation or tightening of the consumer's intertemporal budget constraint can lead both to changes in consumption and to changes in labour supply. Windfall gains represent an unanticipated increase in non-earned income and by reducing an agent's marginal utility of wealth they therefore reduce her incentive to work.

In this chapter we analyze the linkages between windfall gains and working hours using data from the European Community Household Panel Longitudinal Users' Database from 1994 to 2001. We show that an unanticipated rise in wealth reduces working hours in accordance with the life-cycle model, although the effect is, in general, small. The impact of windfall gains is stronger at the external margin, that is, individuals adjust their labour supply primarily by dropping out of the labour force, rather than by reducing their work hours conditional on working.

We also look whether "size matters" with respect to the effects of windfall gains on working hours. We assess how households respond to small, medium or large windfall gains. We find that the effects become stronger as the size of windfall increases. In particular, men receiving a windfall of 50,000 EUR or more, on average reduce labour supply by 1.3 hours per week, which is equivalent to a 3.4% reduction in working hours.

Finally, analysing the effects of windfall gains along various personal characteristics, we find that: (i) at younger and older ages, the effect of windfall gains on labour supply is the most negative; (ii) for married people and people with young children, the windfall gain leads to a stronger decrease in working hours and (iii) for single individuals at the age of around 40, the effect can be positive. A potential explanation for the latter empirical finding is in the effect of windfall gains on reducing liquidity constraints in capital markets. By doing so, windfall gains may encourage people to set up their own business, become self-employed and increase their working hours, as suggested by Lindh and Ohlsson (1996) and Taylor (2001).

This chapter contributes to the literature in following ways. This is the first paper that analyses effects of windfall gains on working hours using European data with more than one country included. Henley (2004) was the first paper done on European data, but he analyzes the effects of capital gains on labour supply using data for Britain only. Furthermore, because we include 15 countries in our analysis, the sample of people for which we observe windfall gains is large, offering a further empirical advantage to our approach. With the panel data set we observe a rich set of personal characteristics of individuals. This gives us an opportunity to better understand the ways that participation and working-hours decisions differ between different types of individuals.

The remainder of the chapter is structured as follows. Section 2 reviews the existing literature on the effects of unexpected variation in income. Section 3 describes the data. Section 4 presents the theoretical and the econometric approach and Section 5 discusses the empirical results. Section 6 concludes.

2 A brief review of the literature

The launch of the pan-European lottery, *Euromillions*, in 2004 induced many people to fantasize about what they would do if they actually won. Notable wins include prizes of around 180 Million EUR which, therefore, reveals the extraordinary importance that a lottery may play in people's life and behaviour.

A vast literature has explored the reaction of consumption and savings to exogenous changes in income. An early example is Bodkin (1959), who used an unexpected National Service Life Insurance dividend paid to veterans of the World War II in 1950. Similarly, Brickman et al. (1978) focused on how the income effect affects consumption. More recent examples include Imbens et al. (2001), who look at the differences among major-prize winners of the Megabucks Lottery in Massachusetts between 1984 and 1988, and Kuhn et al. (2008), who analyze the differences in winnings in the Dutch postcode lottery[1].

Unexpected variation in income may also affect the level of happiness of individuals.[2] Whereas some surveys suggest that money indeed makes people happy (Gardner and Oswald, 2001, 2007), others find only a weak link between unexpected wealth variation and happiness (Myers, 1992; Diener et al., 1999; Argyle, 2001; Nettle, 2005; Layard, 2005).[3]

Another dimension of the effects of exogenous changes in income refers to fiscal policy and, in particular, the effectiveness of temporary fiscal measures. In fact, understanding the effect of unearned income on labour supply is also of great importance for policy makers, as it is at

[1] Some recent studies have also used exogenous variation to analyze neighbourhood and peer effects on individuals (Sacerdote, 2001; Katz et al., 2001; Kling et al., 2005; Ludwig et al., 2001; Kuhn, 2008).
[2] For discussions of this question, see, for example, Easterlin (1974), Martin (1995), and Diener and Biswas-Diener (2002).
[3] Lindahl (2005) shows that higher income from a monetary lottery prize generates good health.

least part of what is needed to evaluate such programs (Kuhn, 2008). For instance, Hankins et al. (2009) show that cash transfers' programs merely postpone bankruptcy of those who are in financial trouble.

In addition to the potential effects of income shocks on consumption and savings or on the level of happiness, a popular belief presumes that the majority of people would quit work if they won a lottery. But do individuals who win continue to work, and if so, why? While the literature on the empirical and theoretical inter-temporal substitution effects in labour supply is well established (Heckman and MaCurdy, 1980; Altonji, 1986), the research on the effects of capital gains is still somewhat insipient (Henley, 2004), despite the fact that lottery winnings are a source of exogenous variation in income (Altonji, 1986).

In the US, Kaplan (1985, 1988) show that the level of education and the type of profession can help explain the percentages of winners who choose to continue to work. Holtz-Eakin et al. (1993) and Imbens et al. (2001) find that windfall gains lead to a reduction in working hours or even a withdrawal from the labour force. In contrast, Joulfaian and Wilhelm (1994) suggest at most a small (although significant) effect for married women and men. Hirschfeld and Eield (2000) use the proposition of work centrality, that is, the degree of importance that working has in one's life at any given time to explain why lotteries may have a limited impact.

In Europe, Blanchflower and Oswald (1998), Taylor (2001), using UK data, and Lindh and Olhsson (1996), based on evidence for Sweden, report a positive effect of windfall gains (inheritance and lottery wins) on the probability of entering self-employment. Arvey et al. (2004) show that the likelihood of quitting work is smaller for individuals who won large amounts in the lottery when they have a greater degree of work centrality. Henley (2004) analyzes the impact of both windfall financial gains and house price shocks on hours worked and suggests that there are significant substitution effects, in particular, in response to house price shocks.

3 Data and descriptive statistics

3.1 Data

The data is obtained from the European Community Household Panel Longitudinal Users' Database (ECHP henceforth). This is a large panel data set that contains household-level and person-level information over time, covering eight survey years from 1994 to 2001. The data includes 15 EU countries: Germany, Denmark, The Netherlands, Belgium, Luxembourg, France, United Kingdom, Ireland, Italy, Greece, Spain, Portugal, Austria, Finland and Sweden. It is an unbalanced panel with a maximum length of 8 years for each individual.

In what follows, the analysis is done at the individual level, rather than at the level of households, with age restricted to 25-60 years. This age band is chosen to avoid complications that arise due to education and retirement choices. The data on incomes and wages are converted using PPP in order to allow for comparisons across countries and over time.

The question of interest relates to the effects of unanticipated windfall gains on labour supply. Working hours are described by the ECHP variable PE005: *Total number of hours working per week (in main + additional jobs)*. In the data, this variable is only available for employed workers. However, we set hours worked to zero for all unemployed individuals and those out of the labour force.

The variable that measures windfall gains is the ECHP variable HF017: *Inherit, receive gift or lottery winnings worth 2000 EURO or more*. It is the response to a following survey question: *"During (... year prior to the survey ...), did anyone in the household inherit any property or capital, or receive a gift or lottery winnings, worth 2000 EURO or more?"*. Observations for which the information on the windfall receipt is missing are discarded.

One major drawback of this variable is that it does not provide information about the exact amount of the windfall gain. However, it can be complemented by the variable HF018: *Amount of the inheritance, gift or lottery winnings*. This variable offers three brackets for the windfall gains: *less than 10,000 EURO, more than 10,000 EURO but less than 50,000 EURO* and *50,000 EURO or more*. We label the three brackets for windfall gains as "small", "medium" and "large", respectively.

These two variables hence give information on the size of windfall gains received by individuals. Nevertheless, given that they are reported in categorical terms, one cannot convert them into PPP terms. As a result, they are not perfectly comparable across countries and over time. Another weakness is that both variables are reported at the household level. Consequently, there is no way to identify which household member was the actual recipient of the windfall gain.[4]

[4] It should be noted, however, that an indicator for a windfall gain is, to some degree, a personal characteristic. For example, in cases where individuals change households (i.e. get married) and they receive windfall gains only after they have moved to a new household, they are recorded as recipients of windfall gains together with their partner. Naturally, individuals from the initial household have not received any windfall gains. Should the

It is important to emphasise that the variable measuring windfall gains is recorded for the "year prior to the survey".[5] Notwithstanding this, we did not decide to adjust the timing of the variable. First, a substantial fraction of the data (that is, 19% of person-year observations) would be lost by lagging the windfall gains variable by one period. Second, leaving the variable as it is, we can be sure that at the time of the interview in the time period *t*, an individual knows whether she has received windfall gains or not. On the contrary, if we lagged windfall gains variable by one period, to *t-1*, we would not know for sure whether at the time of the interview at *t-1* the individual had already received the windfall gains or not[6]. Furthermore, in practice individuals take a bit of time before they react to new economic information. Therefore, it seems more appropriate not to lag the windfall gains variable back by one period.

In Table 1 we report the number of *individuals* in the sample and the number of times they received windfall gains. Only those individuals who were observed at least twice are included. To ease discussion, we label people that have received windfall gains as "winners" and the rest as "non-winners". There are 100,289 individuals in the sample, and most of them (88.4%) never received any inheritance, gift or lottery winnings of more than 2000 EUR. In addition, 8,824 individuals (or a fraction of 8.8%) received windfall gains only once, and about 2% of individuals received windfall gains twice.

For the purpose of the analysis, the most important group is the one with 8,824 individuals who received windfall gains only once, as in the regression analysis it is not straightforward to deal with individuals who received windfall gains more than once. Most of the empirical analysis will therefore be based on that group. Compared to similar research done by other authors, this is quite a large sample and represents one of the advantages of using the ECHP dataset.[7]

individual move households again with a new partner, for example, then he would still be recorded as a recipient of windfall gains, but his new partner would not.

[5] Similarly, income variables are also recorded for "year prior to the survey". On the other hand, net monthly wage and other variables are recorded for "the time of the interview".

[6] How much information the individual possesses at the time of the interview of course depends on the relative timings of windfall gains and survey interview, but on average there is a 50% chance that the individual had already received the windfall gains.

[7] For instance, Imbens et al. (2001) have about 237 winners, Joulfaian and Wilhelm (1994) have 439 heirs in their sample, Holtz-Eakin (1993) have 2,700 married couples and 1632 individuals in their sample, and Henley (2004) has around 5,400 men and women included.

In Table 2 we report the number of individuals by size of windfall gains received. There are 4,172 (48.8%) observed individuals with small windfall gains, 3,353 (39.2%) with medium windfall gains, and 1,023 (12.0%) individuals with large windfall gains.

Table 1: Number of individuals in the sample and number of times they received windfall gains during the period in the sample

# of times windfall gains received	Frequency	Percent
0	88,692	88.44
1	8,824	8.80
2	1,957	1.95
3	501	0.50
4	165	0.16
5	82	0.08
6	26	0.03
7	25	0.02
8	17	0.02
Total	100,289	100.00

Source: European Community Household Panel Longitudinal Users' Database. All individuals of age 25-60. Only individuals who are observed for at least two periods are included.

Table 2: Number of individuals who ever received small/medium/large windfall gains

Size of windfall gains received	Frequency	Percent
small (2000-10,000 EUR)	4,172	48.81
medium (10,000-50,000 EUR)	3,353	39.23
large (more than 50,000 EUR)	1,022	11.96
Total	8,547	100.00

Source: European Community Household Panel Longitudinal Users' Database. All individuals of age 25-60. Only individuals who are observed for at least two periods are included.

3.2 Descriptive statistics

In this sub-section, we analyse differences in personal characteristics between winners and non-winners prior to the receipt of windfall gains, and differences among winners of windfall gains of different sizes (i.e. *small* versus *large* winners). We also compare the means of variables *before* and *after* the receipt of windfall gains.

Table 3 reports the means and number of observations for selected variables, comparing winners, (columns (1) and (2)) and non-winners (columns (3) and (4)). Column (5) reports the *p*-value of the test for differences in means between winners and non-winners. The reported statistics refer to one year *before* the receipt of windfall, which, on average, corresponds to a third year in the sample for winners. Therefore, for non-winners we report the means of the variables in the third year in the sample.

Among the 18 variables reported, only three (the number of children in the household, the percentage of women and the percentage of those who are married) have differences that are not statistically significant. Otherwise, winners tend to be older and they live in slightly smaller households, but for these two variable differences are small. For the rest of the variables, the differences are large and important.

Table 3: Comparing personal information for winners and non-winners (prior to receiving windfall gains)

	have received windfall gains		have not received windfall gains		Diff. in means
	Mean	Obs.	Mean	Obs.	p-value
Variable	(1)	(2)	(3)	(4)	(5)
household size	*3.29*	6,674	*3.46*	75,040	0.000
number of adults (>16) in household	*2.45*	6,674	*2.63*	75,040	0.000
number of children (16<) in household	*0.84*	6,674	*0.83*	75,040	0.739
age	*42.01*	6,674	*41.23*	75,040	0.000
female dummy	*0.51*	6,674	*0.51*	75,040	0.484
married dummy	*0.73*	6,664	*0.72*	74,974	0.249
secondary education dummy	*0.38*	6,576	*0.36*	73,598	0.000
post secondary education dummy	*0.29*	6,576	*0.18*	73,598	0.000
employed dummy	*0.77*	6,656	*0.70*	74,998	0.000
household income	*31,186*	6,648	*25,863*	74,584	0.000
household income - from working	*26,758*	6,651	*21,337*	74,598	0.000
household income - unearned income	*1,448*	6,651	*848*	74,598	0.000
personal income	*15,589*	6,674	*12,095*	75,040	0.000
personal income - from working	*13,270*	6,674	*10,222*	75,040	0.000
personal income - unearned income	*690*	6,674	*376*	75,040	0.000
personal hourly wage	*7.68*	6,674	*6.78*	75,040	0.000
total hours working per week	*30.57*	6,497	*28.26*	73,932	0.000

Source: European Community Household Panel Longitudinal Users' Database. All individuals of age 25-60. Winners are observed one period before receiving windfall gains. This approximately corresponds to period 3 in the sample for non-winners.

Winners are more educated; the share of individuals with post secondary education is 29% for winners and 18% for non-winners; winners are 7 percentage points more likely to be employed than non-winners. According to income variables, winners have higher incomes and wages even *before* windfall gains. By all measures of income (total income, income from working and non-work income), winners are better off than non-winners: the personal total income of winners is about 29% higher and their hourly wage[8] is 13% higher. Higher income

[8] Hourly wage is a measure of offered wages in the labour market. Reported data is in purchasing power parity units in order to be comparable across countries. Hourly wage is calculated from the net monthly wage given in the data, divided by weekly working hours times 4.33 to correct for the average number of weeks in one month.

is partly a consequence of the fact that winners, on average, work more hours per week and they are more likely to be employed. They are also more educated and thus have a higher hourly wage. However, another potential reason for the difference in incomes lays also in the fact that our measure of windfall gains includes gifts and inheritances. It can then be the case that people from better family backgrounds are more likely to receive (large) gifts or inheritances, which is reflected in our data. Family background is of course a fixed effect and will eventually drop out of the analysis when data will be analysed using our econometric methodology[9].

The observed differences between winners and non-winners from Table 3 could of course reflect simply differences across countries. If there were a country with above average number of winners, and also with above average incomes, this would make winners, in a spurious fashion, appear to have higher incomes in the full sample. Data show that in most countries, between 87% and 96% of the sample is comprised of non-winners. However, four countries (Denmark, the Netherlands, Finland and Belgium) have a lower percentage of non-winners, but when we checked differences in means after excluding these four countries, the magnitudes and conclusions were similar. Therefore, we conclude that the differences reported in Table 3 reflect genuine differences between winners and non-winners.

In Table 4, we turn to comparisons of personal characteristics among winners of small, medium and large windfall gains. We report means and number of observations one period prior to the receipt of windfall. Columns (7) – (9) report p-values from testing the null hypothesis of no differences in means between groups.

No statistically significant differences between winners of windfall gains of different sizes are found for household size, number of adults, number of children in household, percentage of females, marital status, and employment status. On the other hand, there are statistically significant differences in age and education: the group with small windfall gains is significantly younger than the other two groups (i.e. 41.4 years compared to 42.5 and 42.8 years for medium and large windfall gains groups, respectively); the group of large winners is also more educated (37% of large winners have education beyond the secondary level, while only 27% of small winners and 28% of medium winners have education of such level).

All hourly wages lower than 1 euro or higher than 100 euros are put to missing. Wages of people who do not work or wages otherwise missing are then imputed. For those individuals for whom wage information is available in some periods but not in others, the average wage of the individual is imputed from the other periods. Other wages are imputed using a regression equation separately for men and women using age, age squared, a married dummy, two education dummies and wave and country dummies as regressors.

[9] This will be discussed in more detail below.

There are also large and highly significant differences in incomes between the three groups; the larger the windfall gains, the higher the income. Such differences in incomes and education can again be explained with family characteristics. If people with higher education and household incomes tend to be from families of better background, then this may be reflected in higher inheritances or gifts. However, this will be controlled for by fixed effects in our estimation.

Table 4: Comparison of personal information among winners by the size of windfall gains received (prior to receiving windfall gains)

	received small windfall gains		received medium windfall gains		received large windfall gains		small vs medium	small vs large	medium vs large
	Mean	Obs.	Mean	Obs.	Mean	Obs.	Difference in means p-value		
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
hh size	*3.30*	3,098	*3.27*	2,579	*3.24*	762	0.358	0.284	0.669
no. of adults in hh	*2.45*	3,098	*2.43*	2,579	*2.44*	762	0.406	0.721	0.852
no. of children in hh	*0.85*	3,098	*0.84*	2,579	*0.80*	762	0.662	0.297	0.470
age	*41.40*	3,098	*42.45*	2,579	*42.77*	762	0.000	0.001	0.428
female dummy	*0.51*	3,098	*0.52*	2,579	*0.52*	762	0.344	0.555	0.974
married dummy	*0.73*	3,095	*0.73*	2,574	*0.71*	761	0.683	0.308	0.207
second. educ. dummy	*0.39*	3,064	*0.37*	2,542	*0.35*	742	0.243	0.029	0.167
post sec. educ. dummy	*0.27*	3,064	*0.28*	2,542	*0.37*	742	0.341	0.000	0.000
employed dummy	*0.77*	3,091	*0.76*	2,573	*0.77*	757	0.463	0.811	0.811
hh income	*28,804*	3,091	*32,185*	2,568	*36,711*	756	0.000	0.000	0.000
hh income - working	*24,958*	3,092	*27,569*	2,570	*31,346*	756	0.000	0.000	0.000
hh income - unearned	*1,386*	3,092	*1,181*	2,570	*2,646*	756	0.077	0.000	0.000
personal income	*14,612*	3,098	*15,987*	2,579	*18,443*	762	0.001	0.000	0.002
pers. income - working	*12,413*	3,098	*13,656*	2,579	*15,717*	762	0.002	0.000	0.006
pers. income - unearned	*665*	3,098	*552*	2,579	*1,280*	762	0.086	0.003	0.000
pers. hourly wage	*7.07*	3,098	*7.98*	2,579	*8.95*	762	0.000	0.000	0.000
weekly hours working	*31.04*	3,016	*30.01*	2,521	*31.22*	736	0.062	0.834	0.166

Source: European Community Household Panel Longitudinal Users' Database. Individuals of age 25-60 that have at some point received windfall gains. Winners are observed one period before receiving windfall gains. Small windfall gains (2000-10,000 EUR), medium windfall gains (10,000-50,000 EUR), large windfall gains (more than 50,000 EUR)

Finally, in Table 5, we compare the means of personal characteristics before and after the receipt of windfall gains. "Before" stands for one period prior to windfall and "after" stands for one period after the windfall. Intuitively, we would expect non-work income to increase from the period before to the period after the receipt of windfall gains. However, this is not necessarily the case, because, strictly speaking, windfall gains bring a one-off increase in non-work income that lasts only for one period. Nevertheless, it is possible that individuals save or invest part of their unanticipated gains and start earning interest, which may increase their non-work income also in subsequent periods. According to the life-cycle theory of labour supply, the receipt of an unexpected windfall should also reduce working hours and employment of the winners.

Table 5: Comparing personal information before and after the receipt of windfall gains

	Before		After		
	Mean	Obs.	Mean	Obs.	Difference in means, p-value
Variable	(1)	(2)	(3)	(4)	(5)
ALL					
employed dummy	*0.77*	6,656	*0.76*	7,079	0.781
hh income	*31,186*	6,648	*31,823*	7,068	0.099
hh income - working	*26,758*	6,651	*27,312*	7,068	0.147
hh income - unearned	*1,448*	6,651	*1,614*	7,068	0.061
personal income	*15,589*	6,674	*15,986*	7,082	0.151
pers. income - working	*13,270*	6,674	*13,592*	7,082	0.228
pers. income - unearned	*690*	6,674	*762*	7,082	0.173
pers. hourly wage	*7.68*	6,674	*7.81*	7.082	0.072
weekly hours working	*30.57*	6,497	*30.56*	6,923	0.966
SMALL WINDFALL GAIN					
employed dummy	*0.77*	3,091	*0.76*	3,444	0.341
hh income	*28,804*	3,091	*28,874*	3,442	0.885
hh income – working	*24,958*	3,092	*24,928*	3,442	0.951
hh income – unearned	*1,386*	3,092	*1,399*	3,442	0.918
personal income	*14,612*	3,098	*14,690*	3,446	0.824
pers. income - working	*12,413*	3,098	*12,478*	3,446	0.847
pers. income - unearned	*665*	3,098	*678*	3,446	0.859
pers. hourly wage	*7.07*	3,098	*7.28*	3,446	0.012
weekly hours working	*31.04*	3,016	*30.57*	3,366	0.362
MEDIUM WINDFALL GAINS					
employed dummy	*0.76*	2,573	*0.78*	2,683	0.261
hh income	*32,185*	2,568	*33,030*	2,675	0.194
hh income - working	*27,569*	2,570	*28,663*	2,675	0.084
hh income - unearned	*1,181*	2,570	*1,430*	2,675	0.035
personal income	*15,987*	2,579	*16,563*	2,684	0.211
pers. income - working	*13,656*	2,579	*14,238*	2,684	0.194
pers. income - unearned	*552*	2,579	*664*	2,684	0.077
pers. hourly wage	*7.98*	2,579	*8.09*	2,684	0.348
weekly hours working	*30.01*	2,521	*30.58*	2,628	0.308
LARGE WINDFALL GAINS					
employed dummy	*0.77*	757	*0.75*	762	0.516
hh income	*36,711*	756	*40,587*	761	0.008
hh income - working	*31,346*	756	*33,858*	761	0.090
hh income - unearned	*2,646*	756	*3,268*	761	0.134
personal income	*18,443*	762	*20,297*	762	0.094
pers. income - working	*15,717*	762	*16,961*	762	0.238
pers. income - unearned	*1,280*	762	*1,497*	762	0.437
pers. hourly wage	*8.95*	762	*9.07*	762	0.691
weekly hours working	*31.22*	736	*31.10*	743	0.912

Source: European Community Household Panel Longitudinal Users' Database. Individuals of age 25-60. Before: one period before receiving windfall gains. After: one period after receiving windfall gains (not in the period when windfall gains were received). Small windfall gains (2000-10,000 EUR), medium windfall gains (10,000-50,000 EUR), large windfall gains (more than 50,000 EUR)

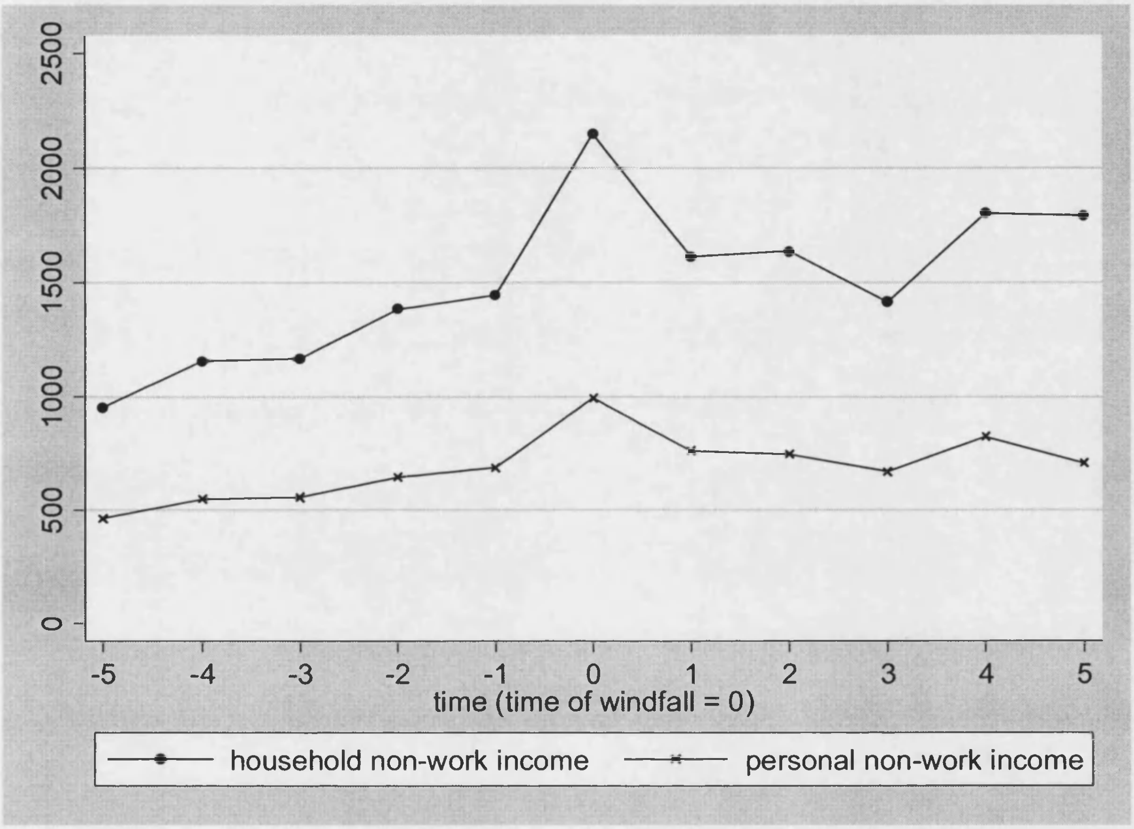
Consider first the top panel of Table 5, where differences for the whole sample are reported. Only three variables are (marginally) significantly different between the two periods: total household income is slightly higher after the receipt of windfall gains at a 10% significance level, household non-work income is higher at a 6% significance level and personal hourly wage is higher at a 7% significance. Weekly hours worked show no difference in the two periods. Looking at the group with small windfall gains, changes in none of the variables are statistically significant from one period to another, except for the hourly wage, which tends to be higher after the receipt of windfall gains. The percentage of employed people and weekly working hours both slightly decrease, but the differences are not significantly different from zero. In the case of individuals who received medium windfall gains, there is a statistically significant rise in the household income from working, in the unearned household income and in the personal unearned income. Interestingly, the share of employed people and weekly working hours show a slight increase, although the differences are not significant. Finally, for the group with large windfall gains, household total income (at a 1% significance level), household income from working (at a 9% significance level) and personal total income (at a 9% significance level) all rise from one period to another. Employment and working hours slightly decrease, but the differences are not statistically significant.

3.3 Non-work income and working hours over time

In this sub-section, we show the evolution of unearned income and working hours over time. From the previous analysis, windfall gains do not seem to have strong effects on income or on labour supply, since differences over time, before and after the windfall gains, are mostly not statistically significant. Hence, one could ask whether the windfall gains variable is a correct measure. For this reason, Figure 1 depicts the average (household and personal) non-work income over time. The time period "0" refers to a time of windfall gains receipt. Since the maximum number of periods for an individual in the sample is eight, the graph is plotted only for five years prior and five years after the receipt of windfall gains. Moving further away from the point of receipt would make the sample size become very small. From Figure 1 it can be seen that the variable windfall gains is meaningful and informative. Indeed there is a positive blip in both household and personal non-work income at the time of receipt. After that, non-work income returns to its upward trend.

Figure 2 displays household income over time by size of windfall gains. Due to limitations in the sample size, we put the large windfall gains and the medium windfall gains groups into a single category. Non-work household income of the medium/large group is, in general, higher than for the small group. The discrete jump in income in the period the windfall gains are received is still visible for both groups, and, as expected, is larger for the group that receives medium/large gains.

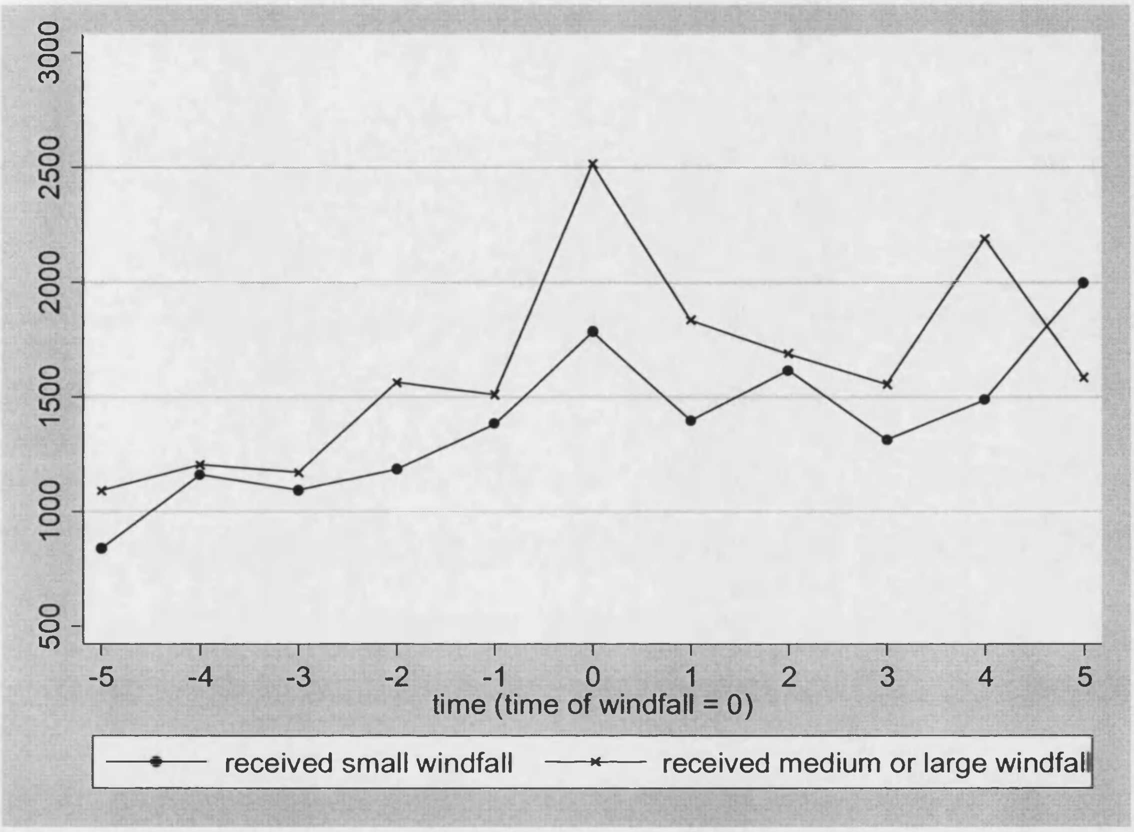
Figure 1: Non-work income over time for those who received windfall gains



Source: European Community Household Panel Longitudinal Users' Database. Individuals who received windfall gains of age 25-60. Time period 0 is period when windfall gains were received.

Next, we turn to the evolution of weekly working hours (Figure 3 and Figure 4). Figure 3 shows that the positive trend in average weekly working hours is reversed after the receipt of windfall gains. Similar information is conveyed by Figure 4, where we split the sample between those who receive small windfall gains and those who receive either medium or large windfall gains. Whereas the evolution of working hours for the small group seems to be more or less unchanged, the downward trend after windfall gains for medium/large group is more apparent. This is consistent with the hypothesis that, after receiving windfall gains, individuals adjust their labour supply downwards. Of course, this is a very crude method of relating working hours to windfall gains and in the analysis that follows we will proceed with the regression analysis.

Figure 2: Non-work household income over time by size of windfall gains



Source: European Community Household Panel Longitudinal Users' Database. Individuals who received windfall gains of age 25-60. Time period 0 is period when windfall gains were received. Small windfall gains (2000-10,000 EUR), medium windfall gains (10,000-50,000 EUR), large windfall gains (more than 50,000 EUR)

4 Theory and econometric approach

4.1 The impact of windfalls on working hours: A theoretical illustration

Consider a representative consumer who chooses consumption, C_t, and leisure hours, L_t, in order to maximize lifetime utility

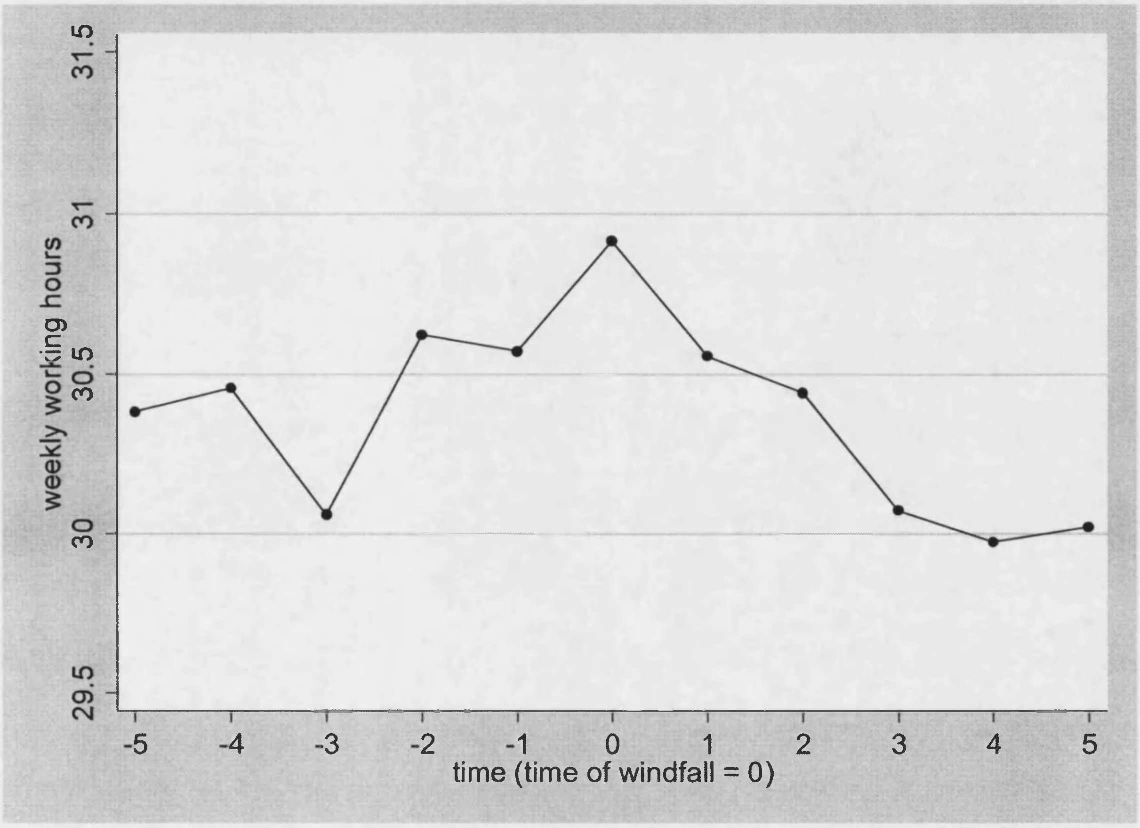
$$\sum_{i=0}^{T}(1+\rho)^{-t}U(C_t, L_t) \tag{1}$$

subject to the intertemporal budget constraint

$$A_0 + \sum_{i=0}^{T} R_t N_t W_t = \sum_{i=0}^{T} R_t C_t \tag{2}$$

where U represents the utility function in time period t that is separable in consumption and leisure, N_t denotes hours worked equal to L^* (a fixed time endowment) minus L_t, A_0 refers to initial assets, W_t is the hourly wage rate, R_t is the discount rate, $\prod_{i=1}^{t} 1/(1+r_i)$, where r is the real rate of interest, and ρ is the rate of time preference.

Figure 3: Working hours per week over time



Source: European Community Household Panel Longitudinal Users' Database. Individuals who received windfall gains of age 25-60. Time period 0 is period when windfall gains were received.

Following MaCurdy (1981), we assume that U has the following form for individual i at time t:

$$U_i(C_{it}, L_{it}) = \alpha_{1it} C_{it}^{\omega_1} - \alpha_{2it} N_{it}^{\omega_2} \tag{3}$$

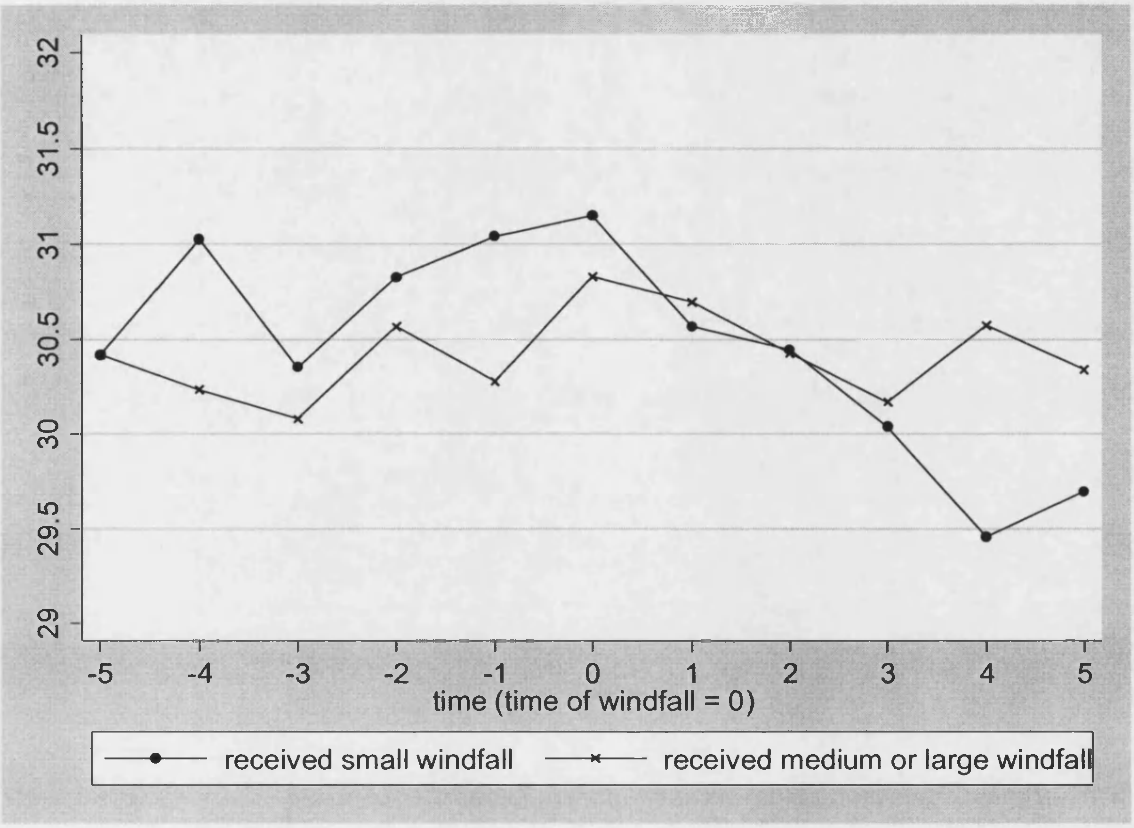
where α_1 and α_2 are 'taste-shifters' which depend on consumer i's preferences at t, $0 < \omega_1 < 1$ and $\omega_2 > 1$.

If we consider an interior optimum (that is, for $N_{it} > 0$), the logarithm of the labour supply function for a given marginal utility of wealth can be expressed as

$$\log N_{it} = (\omega_2 - 1)^{-1}(\log \lambda_{it} - \log \alpha_{2it} - \log \omega_2 + \log(R_t(1+\rho)^t) + \log W_{it}). \quad (4)$$

where λ denotes the marginal utility of wealth.

Figure 4: Working hours per week over time by size of windfall gains



Source: European Community Household Panel Longitudinal Users' Database. Individuals who received windfall gains of age 25-60. Time period 0 is period when windfall gains were received. Small windfall gains (2000-10,000 EUR), medium windfall gains (10,000-50,000 EUR), large windfall gains (more than 50,000 EUR)

Following MaCurdy (1981) we assume that 'tastes' for work are randomly distributed according to the relationship $\log \alpha_{2it} = \gamma X_{it} + \sigma_i + u^*_{it}$ where X_{it} denotes the set of observable determinants of consumer's tastes, σ_i represents the unobserved permanent component of consumer's characteristics and u^*_{it} is a time-varying random component with zero mean.

Assuming a constant real interest rate, replacing the distribution for 'tastes for work' in equation (4) and using the approximation $\log(1+x) \approx x$, we can simplify the labour supply function as:

$$\log N_{it} = -\delta(\sigma_i + \log \omega_2) + \delta(\rho - r)t + \delta \log \lambda_{it} + \delta \log W_{it} + \delta \gamma X_{it} + u_{it} \quad (5)$$

where $\delta = (\omega_2 - 1)^{-1}$, and $u_{it} = \delta u^*_{it}$.

Following Altonji (1986) and Joulfaian and Wilhelm (1994), we assume that the marginal utility of wealth evolves as

$$\log \lambda_{it} = \log \lambda_{it-1} + a + \phi_{it} \tag{6}$$

where ϕ_{it} represents the forecast error of the marginal utility for next period and a is a parameter determined by the discount factor, the interest rates, and the distribution of the forecast error. We approximate λ_{it-1} by

$$\log \lambda_{it-1} = \xi Z_i + \theta \log(E_{t-1}(G_i)) + \varepsilon_i \tag{7}$$

where Z represents the family background characteristics and the effect of the expected lifetime wage profile on the marginal utility, $E_{t-1}[G_i]$ denotes the expected present value of the capital gain (loss), including for example potential inheritance and other windfall gains, and ε_i captures any individual unobserved time invariant heterogeneity in marginal utility of wealth. Combining equations (6) and (7) and plugging into equation (5), we obtain the following labour supply representation:

$$\begin{aligned}\log N_{it} &= \delta(\varepsilon_i - \sigma_i) + \delta\xi Z_i + \delta\theta \log(E_{t-1}(G_i)) - \delta(a + \log \omega_2) + \delta(\rho - r)t + \\ &+ \delta \log W_{it} + \delta\gamma X_{it} + \delta\phi_{it} + u_{it}.\end{aligned} \tag{8}$$

It is clear from the first and the second term on the RHS of (8) that labour supply response should be estimated using fixed effects estimation. Thus one eliminates the need to explicitly control for family background and also removes any potential biases due to ε_i.

When the capital gain is fully unanticipated (that is, $E_{t-1}[G_i]=0$), capital gains affect labour supply *only* via the forecast error, ϕ_{it}. Assuming that the forecast error is a proportion κ of the actual capital gain, that is, $\phi_{it} = \kappa G_{it}$, where $\kappa < 0$, then labour supply response will be $\delta\kappa$, which is negative.

However, when the capital gain is fully anticipated (that is, $E_{t-1}(G_i) = G_i$ and $\phi_{it} = 0$), then capital gains will exert their effects on labour supply by $\delta\theta$. Given that marginal utility would have lowered before the time period in question, there would be no further adjustment at the time of inheritance. Therefore, the unanticipated windfall gains reduce the marginal utility of wealth, and thus reduce labour supply.

4.2 The impact of windfalls on working hours: the econometric specification

Despite the large literature concerned with estimating the impact of unearned income on labour supply, the use of an exogenous measure of income variation is not consensual. As a result, different approaches have been considered, namely: (i) the capital income or spousal-labour earnings as variables measuring unearned income (Imbens et al., 2001); (ii) experimental data with exogenous components of unearned income (Rees, 1974; Pencavel, 1986); and (iii) natural experiments in which large amounts of money were allocated using distribution rules that were independent of preferences and other determinants of economic behaviour (Bodkin, 1959; Kreinin, 1961; Landsberger, 1963; Holtz-Eakin et al., 1993).

We start by looking at whether the windfall gain affects the probability of being employed, and estimate the following linear probability model:

$$\mathrm{Prob}(\mathrm{E}_{\mathrm{it}} = 1) = \mathrm{c}_0 + \mathrm{c}_{0\mathrm{i}} + c_1 Windfall_{it} + c_2 W_{it} + c_3 X_{it} + \varepsilon_{it} \tag{9}$$

for $i = 1, ..., N, t = 1, ..., T$, where E_{it} is a dummy variable that takes the value of 1 if individual i is employed or 0 otherwise, $Windfall_{it}$ is our variable of interest and takes the value of 1 if the household has received a windfall gain or 0 otherwise, W_{it} denotes the hourly wage, X_{it} represents a set of controls for age, civil status and family characteristics, $\mathrm{c}_{0\mathrm{i}}$ is individual fixed effect and ε_{it} is an i.i.d. error term.

In order to assess the effect of unexpected capital gains on working hours, we estimate the empirical counter-part of Equation (8) as described by[10]:

$$\mathrm{H}_{\mathrm{it}} = \mathrm{c}_0 + \mathrm{c}_{0\mathrm{i}} + c_1 Windfall_{it} + c_2 W_{it} + c_3 X_{it} + \varepsilon_{it} \tag{10}$$

for $i = 1, ..., N, t = 1, ..., T$, where H_{it} stands for weekly working hours of household i in year t.

Taking into account that the impact of windfalls on labour supply differs for different amounts of unanticipated gains, we also disaggregate the *Windfall* dummy into three different categories: (i) *Small Windfall*, in the case of capital gains between 2,000 and 10,000 EUR; (ii) *Medium Windfall*, for capital gains between 10,000 and 50,000 EUR; and (iii) *Large Windfall*, when the capital gain exceeds 50,000 EUR. Then, we consider the model:

[10] Additionally assuming $\rho = r$.

$$\mathrm{H}_{it} = \mathrm{c}_0 + \mathrm{c}_{0i} + c_1^1 \mathit{Small\ Windfall}_{it} + c_1^2 \mathit{Medium\ Windfall}_{it} + c_1^3 \mathrm{Large}\ \mathit{Windfall}_{it} + \\ + c_2 W_{it} + c_3 X_{it} + \varepsilon_{it} \tag{11}$$

for $i = 1, ..., N, t = 1, ..., T.$

Finally, we look at whether the effect of the windfall varies with different personal characteristics. Therefore, we interact the regressors with the *Windfall* dummy and estimate the following model:

$$\mathrm{H}_{it} = \mathrm{c}_0 + \mathrm{c}_{0i} + c_1 \mathit{Windfall}_{it} + c_2 W_{it} \times (1 + \mathit{Windfall}_{it}) + c_3 X_{it} \times (1 + \mathit{Windfall}_{it}) + \varepsilon_{it} \tag{12}$$

for $i = 1, ..., N, \ t = 1, ..., T.$

The estimation of the above models is complicated by the potential endogeneity of the wage term on the right-hand side[11]. Altonji (1986) emphasizes that current labour supply depends on all past and expected future wage rates and that it is important to control for permanent differences in wages across individuals.

Consequently, we assess the robustness of the results using both the fixed effects (FE) estimator and the fixed effects instrumental variable (FEIV) estimator. In the IV regressions, we follow Joulfaian and Wilhelm (1994) and Henley (2004) and instrument log monthly wages using conventional earnings-function control variables, namely, two dummies for education, interactions between education dummies and a quadratic in age, as well as country-year dummies.

5 Empirical results

5.1 The effects of windfall gains on working hours

In this and subsequent sub-sections, we analyse the effect of windfall gains on working hours. According to the empirical specification of the theoretical model presented above, we use the fixed effects estimation, thus controlling for family background and other time-invariant personal characteristics. Windfall gains are measured using a dummy variable that takes a value of 1 in the period of windfall receipt and after, and 0 in periods prior to windfall gains. This is in line with the life-cycle model of labour supply where after an unanticipated shock in personal wealth, an individual adjusts her whole labour-supply profile.

[11] Pencavel (1986) also highlights the endogeneity of nonwage income. In the context of our framework we consider windfall gains as unanticipated and exogenous.

In all specifications we include the following set of regressors: the dummy variable for the windfall gain, age, age squared, a dummy for married status and two dummy variables indicating whether there are any children aged 0-6 or 7-15 in the household. We focus on three main specifications, each of them being estimated for the full sample (Table 6), and then separately for men (Table 7) and for women (Table 8).

The first specification (column (1)) analyses the effects of windfall gains on the probability of being employed. According to the theory, after receiving unanticipated windfall gains, individuals are more likely to drop out of the labour force and use windfall gains to enjoy more leisure. Such behaviour is reported in Holtz-Eakin et al. (1993), who find a negative effect of unanticipated inheritances on participation in the labour market. As a dependent variable we use a dummy variable indicating whether or not an individual is employed[12]. Normally, probit specification would be used to analyse this; however, due to inconsistency of probit regression in settings with fixed effects, we use a linear probability model instead[13].

The second specification (column (2)) uses working hours per week as a dependent variable. In this case we include all individuals who ever received windfall gains no matter whether they participate in the labour force in any particular period or not. That is, working hours can take any positive value, but they can also be zero. The second specification thus covers both external and internal margins of adjustment of labour supply to windfall gains[14]. In addition to the standard controls on the right hand side, this specification also includes a measure of hourly wage. As the hourly wage is computed from monthly wages and weekly working hours, it is endogenous. Consequently, we also report the results from an instrumental variables estimation (column (3)). Here we overcome the endogeneity problem by using two dummies for education, interactions of education dummies with quartic in age, as well as country-year dummies as instruments for hourly wage.

Finally, in the third specification (column (4)), we express working hours and wages in logs. As a consequence, only person-year observations with positive working hours and positive wages are included. This specification is closest to the theoretical approach based on the interior solution of the life-cycle optimisation problem derived in section 4.1. From an econometric perspective, however, one should note that a potential problem in this context

[12] In our sample, about 25% of people change their employment status at least once.

[13] In all cases in the paper where we estimated the linear probability model we also checked the results using logit specification with fixed effects. Results were very similar with the same conclusions.

[14] Due to the number of zeroes on the left hand side, we estimated this specification also using the Tobit with fixed effects estimator from Honore (1992). The results were again very similar and conclusions would be the same.

stems from the fact that when receiving windfall gains individuals may decide to reduce their working effort either by reducing working hours *or* by dropping out of the labour force, which can generate a selection bias problem. We present the results both from the fixed effect setting (column (4)) and the instrumental variable estimation with fixed effects (column (5)).

Table 6: Effects of windfall gains on working hours – all

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.00424	-0.0618	0.00194	0.00125	0.00236
	(0.00360)	(0.159)	(0.161)	(0.00365)	(0.00375)
hourly wage		-0.831***	-0.441**		
		(0.0262)	(0.203)		
log hourly wage				-0.404***	-0.197***
				(0.00610)	(0.0469)
age	0.0823***	3.876***	3.625***	0.0564***	0.0386***
	(0.00272)	(0.121)	(0.150)	(0.00295)	(0.00464)
age^2	-0.000986***	-0.0452***	-0.0439***	-0.000490***	-0.000391***
	(3.09e-05)	(0.00137)	(0.00142)	(3.40e-05)	(3.88e-05)
married dummy	0.0129*	1.205***	1.157***	0.0146**	0.00932
	(0.00680)	(0.301)	(0.312)	(0.00671)	(0.00708)
children 0-6 in household dummy	-0.0516***	-2.841***	-2.961***	-0.0491***	-0.0534***
	(0.00571)	(0.253)	(0.259)	(0.00560)	(0.00584)
children 7-15 in household dummy	-0.0425***	-2.318***	-2.333***	-0.0293***	-0.0292***
	(0.00519)	(0.230)	(0.233)	(0.00518)	(0.00533)
Constant	-0.841***	-41.45***	-36.29***	2.974***	3.131***
	(0.0589)	(2.615)	(3.089)	(0.0617)	(0.0680)
Observations	54164	53011	52348	40239	39789
Number of individuals	10395	10357	10322	8735	8700
R-squared	0.024	0.049		0.126	
Hausman test[+]			0.000		0.000

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. [+] Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

Table 6 summarizes the findings for the effects of windfall gains on working hours using the whole sample. The empirical evidence is not supportive of theoretical prediction of the life-cycle model that an unanticipated rise in wealth reduces an individual's working hours via a reduction in the marginal utility of wealth, in particular, the coefficients of the windfall gains dummy are not statistically significant. On the other hand, most coefficients of the other control variables have the expected signs and magnitudes and are statistically significant: (i) age has a nonlinear, inverted "U" shaped effect on labour supply; (ii) being married tends to increase labour supply; and (iii) having children reduces it. Only hourly wage has a surprising negative effect on working hours, which is reduced after using instrumental variables estimation, but it nevertheless remains negative and significant[15].

Table 7: Effects of windfall gains on working hours – men

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.00493	-0.158	-0.0525	0.00126	0.00130
	(0.00443)	(0.237)	(0.239)	(0.00398)	(0.00408)
hourly wage		-0.879***	-0.812***		
		(0.0360)	(0.218)		
log hourly wage				-0.368***	-0.192***
				(0.00704)	(0.0462)
age	0.0955***	5.188***	5.064***	0.0621***	0.0471***
	(0.00338)	(0.181)	(0.205)	(0.00323)	(0.00490)
age^2	-0.00115***	-0.0602***	-0.0593***	-0.000580***	-0.000495***
	(3.83e-05)	(0.00205)	(0.00208)	(3.69e-05)	(4.23e-05)
married dummy	0.0197**	2.126***	2.050***	0.0349***	0.0293***
	(0.00843)	(0.449)	(0.460)	(0.00748)	(0.00785)
children 0-6 in household dummy	-0.0309***	-1.207***	-1.406***	3.02e-05	-0.00319
	(0.00684)	(0.366)	(0.372)	(0.00602)	(0.00625)
children 7-15 in household dummy	-0.0356***	-1.497***	-1.579***	-0.00605	-0.00580
	(0.00640)	(0.341)	(0.346)	(0.00566)	(0.00582)
Constant	-0.980***	-59.98***	-56.94***	2.959***	3.076***
	(0.0733)	(3.925)	(4.335)	(0.0679)	(0.0738)
Observations	26176	25626	25327	22503	22278
Number of individuals	5087	5073	5061	4691	4679
R-squared	0.044	0.070		0.140	
Hausman test[+]			0.000		0.008

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. [+] Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

Table 7 reports the results for the sample consisting of men. Again, the coefficients of the windfall gains dummy are not statistically significant. Nevertheless, they are negative, which is in accordance with the theoretical formulation of the life-cycle model. Other controls have coefficients of sensible signs and magnitudes and, in particular, one can see that being married increases men's labour supply.

Table 8 shows the results for the sample consisting of women. Again, in all five specifications, the effects of windfall gains are not statistically significant and are close to zero. These findings are therefore in line with Imbens et al. (2001), who show that the

[15] This is perhaps due to the income effect being stronger than the substitution effect, although empirically this remains an unresolved issue. All regressions reported in the chapter have also been run using net monthly wage instead. In this case, the coefficient on the wage variable was always positive in both the OLS and instrumental variables specifications. However, the coefficients on the other variables and the conclusions regarding the effects of windfall gains were very similar to the results from the specifications which included the hourly wage.

reaction of people to non-earned income does not differ significantly between men and women.

Table 8: Effects of windfall gains on working hours – women

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.00367	0.0201	0.0563	0.000261	0.00366
	(0.00558)	(0.213)	(0.216)	(0.00659)	(0.00687)
hourly wage		-0.759***	-0.0268		
		(0.0385)	(0.330)		
log hourly wage				-0.441***	-0.149**
				(0.0103)	(0.0752)
age	0.0699***	2.651***	2.259***	0.0474***	0.0230***
	(0.00419)	(0.160)	(0.216)	(0.00532)	(0.00763)
age^2	-0.000832***	-0.0314***	-0.0294***	-0.000357***	-0.000232***
	(4.77e-05)	(0.00182)	(0.00195)	(6.21e-05)	(6.81e-05)
married dummy	0.00581	0.179	0.190	-0.0135	-0.0198
	(0.0105)	(0.402)	(0.420)	(0.0118)	(0.0125)
children 0-6 in household dummy	-0.0757***	-4.828***	-4.893***	-0.122***	-0.130***
	(0.00915)	(0.350)	(0.362)	(0.0103)	(0.0109)
children 7-15 in household dummy	-0.0512***	-3.309***	-3.232***	-0.0635***	-0.0643***
	(0.00808)	(0.308)	(0.313)	(0.00932)	(0.00974)
Constant	-0.705***	-23.60***	-16.21***	3.015***	3.245***
	(0.0907)	(3.472)	(4.363)	(0.110)	(0.121)
Observations	27988	27385	27021	17736	17511
Number of individuals	5308	5284	5261	4044	4021
R-squared	0.015	0.036		0.128	
Hausman test+			0.003		0.001

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

So far we have found no supportive evidence that the receipt of unanticipated windfall gains has a significant and negative impact on labour supply. This finding however could either reflect that these effects are non-existent, or that they are very small and not well captured in the data. In fact, other researchers have reported that unanticipated windfall gains (inheritances, financial wealth, housing prices) have, at most, a small impact on labour supply, which in many cases are only marginally statistically significant (Holtz-Eakin, 1993; Joulfaian and Wilhelm, 1994; Henley, 2004). Another possible explanation for the small and not statistically significant results may lay in the fact that the windfall gains dummy used so far does not contain enough information. This is because it simply indicates whether an individual has received a windfall gain or not, no matter what the size of the gain was. Therefore, in the next sub-section, we split the windfall dummy into three groups: small, medium and large.

5.2 The effects of small, medium and large windfall gains

We now introduce into the regression three dummies representing the size of windfall gains that individuals receive: small (less than 10,000 EURO), medium (more than 10,000 EURO but less than 50,000 EURO) and large (50,000 EURO or more). The benchmark for comparison is the time before the windfall gains are received. For example, the small windfall gains dummy tells us by how much working hours decrease (or increase), on average, due to the windfall gain in comparison to the situation where the windfall gain has not yet been received. According to the theory, the higher the unanticipated windfall gain, the stronger the effect on the marginal utility of wealth and the more negative we expect the effect on labour supply to be. Therefore, we expect the effect of the large windfall gains dummy to be negative and largest in absolute value.

Table 9: Effects of small/medium/large windfall gains on working hours – all

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
small windfall gains dummy	-0.00283	-0.132	-0.114	0.00671	0.00599
	(0.00458)	(0.203)	(0.204)	(0.00464)	(0.00476)
medium windfall gains dummy	-0.00597	-0.0102	0.0740	-0.00356	-0.00143
	(0.00494)	(0.218)	(0.221)	(0.00502)	(0.00517)
large windfall gains dummy	-0.0213**	-0.568	-0.406	-0.0137	-0.00931
	(0.00835)	(0.371)	(0.378)	(0.00851)	(0.00882)
hourly wage		-0.830***	-0.427**		
		(0.0262)	(0.204)		
log hourly wage				-0.404***	-0.200***
				(0.00610)	(0.0471)
age	0.0824***	3.885***	3.630***	0.0565***	0.0388***
	(0.00272)	(0.121)	(0.150)	(0.00295)	(0.00464)
age^2	-0.000985***	-0.0452***	-0.0439***	-0.000488***	-0.000391***
	(3.09e-05)	(0.00137)	(0.00142)	(3.40e-05)	(3.88e-05)
married dummy	0.0131*	1.213***	1.159***	0.0148**	0.00942
	(0.00680)	(0.301)	(0.312)	(0.00671)	(0.00708)
children 0-6 in household dummy	-0.0518***	-2.848***	-2.970***	-0.0491***	-0.0534***
	(0.00571)	(0.253)	(0.259)	(0.00560)	(0.00584)
children 7-15 in household dummy	-0.0427***	-2.323***	-2.338***	-0.0294***	-0.0293***
	(0.00519)	(0.230)	(0.233)	(0.00518)	(0.00533)
Constant	-0.848***	-41.78***	-36.56***	2.971***	3.126***
	(0.0587)	(2.609)	(3.085)	(0.0616)	(0.0679)
Observations	54164	53011	52348	40239	39789
Number of individuals	10395	10357	10322	8735	8700
R-squared	0.024	0.049		0.127	
Hausman test+			0.000		0.000

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

We report the results separately for the whole sample, for men and for women, in Table 9, Table 10 and Table 11, respectively.

Column (1) of Table 9 suggests that windfall gains have a negative effect on the probability of being employed. Despite not being statistically significant for small and medium windfall gains, the coefficient associated with large windfall gains is negative and statistically significant. This lends some empirical support to the idea that after receiving large windfall gains, individuals are more likely to drop out of the labour force. As for the other specifications (columns (2) - (5)), the results are not statistically significant, although large windfall gains tend to have the most negative impact on labour supply.

Table 10: Effects of small/medium/large windfall gains on working hours – men

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
small windfall gains dummy	-0.00104	-0.177	-0.126	0.00557	0.00517
	(0.00563)	(0.301)	(0.303)	(0.00502)	(0.00515)
medium windfall gains dummy	-0.00617	0.0221	0.111	0.00139	0.00153
	(0.00610)	(0.326)	(0.329)	(0.00552)	(0.00566)
large windfall gains dummy	-0.0306***	-1.337**	-0.966*	-0.0208**	-0.0184*
	(0.0103)	(0.552)	(0.560)	(0.00927)	(0.00953)
hourly wage		-0.878***	-0.795***		
		(0.0360)	(0.219)		
log hourly wage				-0.368***	-0.196***
				(0.00704)	(0.0463)
age	0.0954***	5.191***	5.065***	0.0620***	0.0473***
	(0.00338)	(0.181)	(0.206)	(0.00323)	(0.00490)
age^2	-0.00115***	-0.0601***	-0.0592***	-0.000577***	-0.000495***
	(3.84e-05)	(0.00205)	(0.00209)	(3.69e-05)	(4.23e-05)
married dummy	0.0199**	2.138***	2.055***	0.0350***	0.0295***
	(0.00843)	(0.449)	(0.460)	(0.00748)	(0.00784)
children 0-6 in household dummy	-0.0311***	-1.221***	-1.418***	-6.81e-05	-0.00316
	(0.00684)	(0.366)	(0.372)	(0.00602)	(0.00625)
children 7-15 in household dummy	-0.0357***	-1.505***	-1.586***	-0.00606	-0.00580
	(0.00640)	(0.341)	(0.346)	(0.00566)	(0.00582)
Constant	-0.983***	-60.20***	-57.13***	2.960***	3.075***
	(0.0730)	(3.914)	(4.333)	(0.0677)	(0.0737)
Observations	26176	25626	25327	22503	22278
Number of individuals	5087	5073	5061	4691	4679
R-squared	0.044	0.070		0.141	
Hausman test+			0.000		0.034

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

Table 10 displays the results for the sample of men. In all five specifications, not only is the coefficient of large windfall gains the most negative, but it is also statistically significant in all cases. Column (1) indicates that receiving large unanticipated windfall gains induces some men to leave the labour force. Similarly, in specifications with working hours on the left hand side (columns (2) – (5)) there is a statistically significant and negative effect of large windfall gains on working hours.

Column (2) captures both external and internal margins of adjustment: on average, receiving an unanticipated windfall gain of 50,000 EUR or more reduces the labour supply of men by 1.3 hours per week. Since average working hours for men in the sample are equal to 39.2 hours per week, this represents on average a 3.4% reduction in working hours. In column (4), where only the internal margin is considered, the evidence suggests that the large windfall gains reduce working hours by 2.1%. Since only the adjustment of working hours conditional on working is taken into account, the effect is plausibly smaller.

Table 11: Effects of small/medium/large windfall gains on working hours – women

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
small windfall gains dummy	-0.00442	-0.0920	-0.127	0.00806	0.00690
	(0.00712)	(0.272)	(0.275)	(0.00845)	(0.00877)
medium windfall gains dummy	-0.00583	-0.0449	0.0757	-0.0111	-0.00479
	(0.00765)	(0.291)	(0.297)	(0.00897)	(0.00943)
large windfall gains dummy	-0.0123	0.211	0.183	-0.00496	0.00332
	(0.0129)	(0.495)	(0.508)	(0.0154)	(0.0163)
hourly wage		-0.759***	-0.0251		
		(0.0385)	(0.330)		
log hourly wage				-0.442***	-0.148**
				(0.0103)	(0.0754)
age	0.0702***	2.660***	2.270***	0.0476***	0.0231***
	(0.00419)	(0.160)	(0.215)	(0.00532)	(0.00763)
age^2	-0.000833***	-0.0315***	-0.0294***	-0.000357***	-0.000231***
	(4.77e-05)	(0.00182)	(0.00195)	(6.21e-05)	(6.81e-05)
married dummy	0.00593	0.178	0.192	-0.0135	-0.0199
	(0.0105)	(0.402)	(0.419)	(0.0118)	(0.0126)
children 0-6 in household dummy	-0.0758***	-4.824***	-4.893***	-0.122***	-0.130***
	(0.00915)	(0.350)	(0.363)	(0.0103)	(0.0109)
children 7-15 in household dummy	-0.0513***	-3.308***	-3.232***	-0.0636***	-0.0643***
	(0.00808)	(0.308)	(0.313)	(0.00932)	(0.00974)
Constant	-0.715***	-23.90***	-16.58***	3.009***	3.237***
	(0.0906)	(3.466)	(4.341)	(0.110)	(0.121)
Observations	27988	27385	27021	17736	17511
Number of individuals	5308	5284	5261	4044	4021
R-squared	0.015	0.036		0.129	
Hausman test+			0.005		0.004

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

The results for women in Table 11, on the other hand, are not supportive of the idea that windfall gains reduce labour supply. None of the coefficients on windfall gains dummies are statistically significant, and many of them are nonnegative. In light of the fact that women are usually considered as being less attached to the labour market than men, this result is surprising. One would indeed expect the effect of windfall gains to be stronger and more negative for women. For instance, Henley (2004) finds a significant adjustment in hours worked to unanticipated financial gains for both men and women, but the largest impact occurs for women. In addition, the author shows that while men seem to make a (positive) adjustment in hours to housing losses but not a (negative) adjustment to gains, for women the reverse is true. On the other hand, Imbens et al. (2001) estimate the marginal propensity to earn out of unearned income and find that it does not differ significantly between men and women. Joulfaian and Wilhelm (1994) also show that the hours' reduction by married women is of the same order of magnitude as men's.

A potential reason why the effects do not come through very strongly in the analysis above is due to heterogeneity in the effects of windfall gains. The effects of windfall gains may possibly depend on personal characteristics in an important way, so that the impact cancels out across different individuals. For example, an old worker may retire from the labour market after receiving the windfall gain, whereas a young employed worker may use it as a starting capital for a new business, become self-employed and increase his labour supply[16]. In order to assess the potential heterogeneity in the effects of windfall gains, in the next sub-section we interact the windfall gains dummy with several personal characteristics.

5.3 The effects of windfall gains and interactions

In order to see whether the effects of unanticipated positive wealth shocks differ with respect to personal characteristics, we interact the windfall gains dummy with the other regressors. As discussed above, if the effects of windfall gains depend on personal characteristics and if, in addition, they cancel each other out across individuals, then this would explain why in the previous regressions the effects were small and often not statistically significant.

In Table 12 we report the results for the whole sample. Note that the coefficient in front of the windfall gain dummy alone simply represents the effect of windfall gains when all other controls are set to zero. In most cases, the coefficients of the interaction terms are statistically significant. Column (1) displays the results for the probability of being employed. Looking at the interactions, age has a "U" shaped curve effect: at younger and older ages, the effect of

[16] Lindh and Ohlsson (1996) and Taylor (2001) find empirical evidence of such behaviour. The authors argue that windfall gains can contribute to the relaxation of the liquidity constraints in capital markets and they find evidence that windfall gains increase the probability of becoming self-employed.

windfall gains on participation is the most negative. This is in line with evidence reported by Holtz-Eakin (1993) and Imbens et al. (2001).

Table 12: Effects of windfall gains and interactions – all

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.351***	-11.64***	-11.52***	0.0568	-6.97e-05
	(0.0582)	(2.579)	(2.633)	(0.0631)	(0.0670)
hourly wage		-0.912***	-0.158		
		(0.0334)	(0.300)		
interaction of windfall gains with		0.117***	-0.187		
wage		(0.0287)	(0.118)		
log hourly wage				-0.399***	-0.193***
				(0.00714)	(0.0475)
interaction of windfall gains with				-0.00797	0.00363
log wage				(0.00617)	(0.0196)
age	0.0700***	3.371***	2.969***	0.0529***	0.0324***
	(0.00378)	(0.168)	(0.207)	(0.00410)	(0.00569)
interaction of windfall gains with	0.0186***	0.637***	0.726***	-0.000138	0.00192
age	(0.00291)	(0.129)	(0.134)	(0.00318)	(0.00339)
age^2	-0.000843***	-0.0390***	-0.0367***	-0.000440***	-0.000312***
	(4.53e-05)	(0.00201)	(0.00209)	(4.97e-05)	(5.52e-05)
interaction of windfall gains with	-0.000217***	-0.00792***	-0.00869***	-7.41e-06	-3.52e-05
age^2	(3.45e-05)	(0.00153)	(0.00157)	(3.82e-05)	(4.02e-05)
married dummy	0.0398***	2.082***	2.115***	0.0268***	0.0221***
	(0.00790)	(0.350)	(0.361)	(0.00781)	(0.00824)
interaction of windfall gains with	-0.0436***	-1.505***	-1.619***	-0.0233***	-0.0244***
married dummy	(0.00694)	(0.308)	(0.316)	(0.00702)	(0.00724)
children 0-6 in hhold dummy	-0.0425***	-1.978***	-2.251***	-0.0332***	-0.0371***
	(0.00740)	(0.328)	(0.336)	(0.00732)	(0.00760)
interaction of windfall gains with	-0.00801	-1.223***	-0.947***	-0.0247***	-0.0247***
children 0-6 dummy	(0.00783)	(0.347)	(0.358)	(0.00783)	(0.00827)
children 7-15 in hhold dummy	-0.0433***	-2.135***	-2.149***	-0.0303***	-0.0312***
	(0.00676)	(0.299)	(0.304)	(0.00679)	(0.00699)
interaction of windfall gains with	0.00822	0.0116	0.0738	0.00755	0.00931
children 7-15 dummy	(0.00755)	(0.334)	(0.339)	(0.00756)	(0.00778)
Constant	-0.610***	-32.06***	-25.22***	3.009***	3.222***
	(0.0765)	(3.395)	(3.928)	(0.0807)	(0.0897)
Observations	54164	53011	52348	40239	39789
Number of individuals	10395	10357	10322	8735	8700
R-squared	0.026	0.051		0.128	
Hausman test+			0.001		0.000

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

Being married also makes the negative effect of windfall gains on participation stronger. Thus, married people are more likely to reduce their labour supply. One possible explanation for this is that married people have less to worry about their income situation, as they have a partner who usually earns an income. Such behaviour can also be explained with a social multiplier (Glaeser et al., 2003). If people like to spend leisure in groups, then they are more likely to consume leisure when they have someone to spend the leisure with.

Table 13: Effects of windfall gains and interactions – men

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.292***	-12.85***	-12.35***	-0.00101	-0.0175
	(0.0719)	(3.857)	(3.929)	(0.0688)	(0.0718)
hourly wage		-0.933***	-0.510		
		(0.0446)	(0.334)		
interaction of windfall gains with wage		0.0820**	-0.212		
		(0.0393)	(0.144)		
log hourly wage				-0.366***	-0.185***
				(0.00810)	(0.0465)
interaction of windfall gains with log wage				-0.00325	-0.0191
				(0.00685)	(0.0201)
age	0.0862***	4.614***	4.361***	0.0560***	0.0388***
	(0.00470)	(0.252)	(0.288)	(0.00448)	(0.00612)
interaction of windfall gains with age	0.0154***	0.687***	0.750***	0.00231	0.00469
	(0.00360)	(0.193)	(0.200)	(0.00347)	(0.00373)
age^2	-0.00104***	-0.0531***	-0.0516***	-0.000498***	-0.000386***
	(5.60e-05)	(0.00300)	(0.00307)	(5.39e-05)	(6.03e-05)
interaction of windfall gains with age^2	-0.000180***	-0.00862***	-0.00898***	-3.98e-05	-6.76e-05
	(4.26e-05)	(0.00228)	(0.00233)	(4.14e-05)	(4.38e-05)
married dummy	0.0353***	2.579***	2.514***	0.0387***	0.0344***
	(0.00992)	(0.529)	(0.542)	(0.00884)	(0.00924)
interaction of windfall gains with	-0.0251***	-0.735	-0.758	-0.00908	-0.0100
married dummy	(0.00904)	(0.485)	(0.494)	(0.00824)	(0.00850)
children 0-6 in hhold dummy	-0.0152*	-0.328	-0.726	0.0130*	0.00975
	(0.00897)	(0.479)	(0.488)	(0.00790)	(0.00817)
interaction of windfall gains with	-0.0214**	-1.235**	-0.893*	-0.0224***	-0.0210**
children 0-6 dummy	(0.00955)	(0.510)	(0.519)	(0.00845)	(0.00881)
children 7-15 in hhold dummy	-0.0321***	-1.285***	-1.408***	0.00313	0.00272
	(0.00839)	(0.448)	(0.455)	(0.00749)	(0.00770)
interaction of windfall gains with	-0.000463	-0.112	0.0145	-0.0125	-0.0105
children 7-15 dummy	(0.00939)	(0.502)	(0.507)	(0.00834)	(0.00855)
Constant	-0.807***	-49.14***	-44.71***	3.053***	3.198***
	(0.0951)	(5.093)	(5.583)	(0.0887)	(0.0976)
Observations	26176	25626	25327	22503	22278
Number of individuals	5087	5073	5061	4691	4679
R-squared	0.045	0.072		0.141	
Hausman test+			0.001		0.048

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

In Columns (2) and (3) where total working hours are used as a dependent variable, results are similar: (i) there is an indication of a non-linear effect of age; (ii) being married makes the effect of windfall gains more negative; and (iii) similarly, having children of age 0-6 leads to a larger decrease in working hours, indicating that parents of young children do not drop out of the labour force after receiving windfall gains, but they seek shorter working hours. Columns (4) and (5) show the results for the sample with positive hours worked. They confirm that being married and having small children makes the effect of windfall gains more negative.

Table 13 reports the empirical findings for the sample of men. In Columns (1), (2) and (3), one can see that if they are of younger or older age, have young children or are married, then windfall gains reduce the probability of being employed. However, when we take away the adjustment along the external margin (columns (4) and (5)), the effects largely lose statistical significance, except for the effects of having young children.

Table 14 provides similar findings for women. In Column (1), where the probability of being employed is explored, there is a "U" shaped effect of age and a statistically significant negative effect of being married. The coefficients on the interactions of children dummies with windfall gains are not statistically significant. In the regressions with working hours on the left hand side (columns (2) and (3)), the results show that women with young children tend to reduce working hours after the receipt of windfall gains. Columns (4) and (5) also corroborate this finding.

In order to obtain a better understanding of the heterogeneity in the effects of windfall gains and how they depend on personal characteristics, Table 15 reports the predicted effects of windfall gains on labour supply by various types of individuals. We compute the effects of windfall gains for 8 hypothetical types of individuals. We start by choosing four different age groups: 25, 30, 40 or 55 years. We assume that: (i) individuals of age 25 can only be single and have no children; (ii) individuals of age 30 are either single with no children, or married with one child of age 0-6, (iii) individuals of age 40 can be single without children, married without children, or married with one child of age 0-6 and one child of age 7-15; (iv) individuals of age 55 can either be single or married, but they do not have any young children.

Table 14: Effects of windfall gains and interactions – women

Dependent variable:	Employment dummy	Working hours per week		Log working hours per week	
	(1)	(2)	(3)	(4)	(5)
	FE	FE	FE IV	FE	FE IV
windfall gains dummy	-0.387***	-7.540**	-8.464**	0.213*	0.0951
	(0.0909)	(3.465)	(3.528)	(0.115)	(0.125)
hourly wage		-0.919***	-0.224		
		(0.0512)	(0.425)		
interaction of windfall gains with wage		0.210***	0.115		
		(0.0442)	(0.177)		
log hourly wage				-0.438***	-0.170**
				(0.0124)	(0.0786)
interaction of windfall gains with log wage				-0.00632	0.0393
				(0.0113)	(0.0383)
age	0.0558***	2.361***	1.944***	0.0511***	0.0246***
	(0.00586)	(0.224)	(0.283)	(0.00741)	(0.00950)
interaction of windfall gains with age	0.0204***	0.401**	0.482***	-0.00848	-0.00642
	(0.00453)	(0.173)	(0.182)	(0.00581)	(0.00626)
age^2	-0.000670***	-0.0278***	-0.0251***	-0.000394***	-0.000246**
	(7.03e-05)	(0.00268)	(0.00284)	(9.10e-05)	(9.82e-05)
interaction of windfall gains with age^2	-0.000236***	-0.00500**	-0.00596***	9.49e-05	6.62e-05
	(5.41e-05)	(0.00206)	(0.00214)	(7.04e-05)	(7.51e-05)
married dummy	0.0403***	1.212***	1.279***	0.00334	-0.00508
	(0.0121)	(0.464)	(0.482)	(0.0135)	(0.0145)
interaction of windfall gains with married dummy	-0.0549***	-1.788***	-1.803***	-0.0315***	-0.0288**
	(0.0103)	(0.396)	(0.414)	(0.0117)	(0.0125)
children 0-6 in hhold dummy	-0.0739***	-4.075***	-4.127***	-0.100***	-0.107***
	(0.0117)	(0.448)	(0.460)	(0.0134)	(0.0141)
interaction of windfall gains with children 0-6 dummy	0.00775	-1.070**	-1.017**	-0.0330**	-0.0340**
	(0.0124)	(0.475)	(0.492)	(0.0144)	(0.0154)
children 7-15 in hhold dummy	-0.0564***	-3.356***	-3.230***	-0.0778***	-0.0807***
	(0.0105)	(0.400)	(0.406)	(0.0121)	(0.0127)
interaction of windfall gains with children 7-15 dummy	0.0176	0.406	0.353	0.0333**	0.0353**
	(0.0117)	(0.447)	(0.454)	(0.0135)	(0.0141)
Constant	-0.440***	-18.02***	-10.49**	2.912***	3.231***
	(0.118)	(4.520)	(5.328)	(0.145)	(0.159)
Observations	27988	27385	27021	17736	17511
Number of individuals	5308	5284	5261	4044	4021
R-squared	0.017	0.039		0.130	
Hausman test+			0.007		0.007

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. + Reports p-value of the Hausman test of endogeneity. In IV regression (log) monthly wage is instrumented using two dummies for education, interactions between education dummies and quartic in age and country-year dummies.

In those specifications where hourly wage is one of the regressors, predicted effects are computed at the mean wage in the sample. Specification (1) represents effects from participation regressions, specification (2) reports the FE IV regressions with working hours per week on the left hand side[17], and specification (3) reports effects from FE IV regressions with log working hours as the dependent variable. In brackets we also report p-values from the Wald test of whether the computed predicted effects are statistically different from zero.

For young single people who have no children (first row of Table 15), receiving windfall gains reduces the probability of participating in the labour force. This can be seen, for example, in column (1), where the coefficient is -0.021. A possible explanation for this finding is that after winning an unanticipated windfall young people may decide to prolong their education and finance it with the money they have won. However, as specification (3) suggests, conditional on being employed, windfall gains *increase* the labour supply of young people, perhaps inducing young people to become self-employed and work more.

For individuals of age 30 and with no children, there is a positive effect of windfall gains on the probability of participating in the labour market and also on working hours conditional on working. On the other hand, for individuals of the same age who are married and have children of young age, the effects of windfall gains on labour supply are negative, with effects being strongly negative for both men and women.

In the case of individuals aged 40 with no spouse and no children, there is a *positive* effect of windfall gains on labour supply. Weekly working hours, on average, increase by 2.152 hours (specification (2)) or, conditional on working, for 0.028 log hours (specification (3)). This piece of evidence can again be related with the rise in probability of becoming self-employed as discussed in Lindh and Ohlsson (1996) and Taylor (2001). Interestingly, however, these effects decrease and end up being not statistically significant for individuals of age 40 who are married, especially if they have children.

We also find support in our data for the behaviour suggested by Lindh and Ohlsson (1996) and Taylor (2001). The results are reported in the Appendix. We find that winning windfall gains increases the probability of becoming self-employed for men. Including interaction terms supports the story that men aged about 40 years have a higher probability of becoming self-employed after winning a windfall gain. However, married men have a higher probability of becoming self-employed compared to single men.

Turning back to the results in Table 15, the effects of windfall gains on the labour supply of married individuals aged 55 tend to be statistically significant and negative. The impact of

[17] The results of this specification were also checked using the Honore (1992) estimator, without instruments. The results were very similar and conclusions would be unchanged.

windfall gains in the employment participation (specification (1)) is equal to -0.028 and the effect on weekly working hours (specification (2)) is equal to -0.966. These general patterns are similar for men and for women.

Table 15: Effects of windfall gains on labour supply by types of individuals

	Full sample			Men			Women		
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
25 yrs, single, no children	-0.021**	-0.259	0.033***	-0.020*	-1.002	0.019	-0.025	0.682	0.049**
	(0.034)	(0.609)	(0.005)	(0.091)	(0.182)	(0.173)	(0.119)	(0.311)	(0.016)
30 yrs, single, no children	0.012*	0.979***	0.033***	0.007	0.279	0.024**	0.012	1.455***	0.035**
	(0.093)	(0.006)	(0.000)	(0.411)	(0.597)	(0.011)	(0.318)	(0.002)	(0.011)
30 yrs, married, with one child	-0.039***	-1.587***	-0.016**	-0.040***	-1.372***	-0.007	-0.035***	-1.365***	-0.028*
	(0.000)	(0.000)	(0.041)	(0.000)	(0.010)	(0.410)	(0.002)	(0.003)	(0.054)
40 yrs, single, no children	0.046***	2.152***	0.028***	0.034***	1.494***	0.023***	0.050***	2.107***	0.017
	(0.000)	(0.000)	(0.000)	(0.000)	(0.003)	(0.007)	(0.000)	(0.000)	(0.202)
40 yrs, married, no children	0.003	0.533*	0.003	0.009	0.736	0.013*	-0.005	0.305	-0.011
	(0.701)	(0.083)	(0.650)	(0.275)	(0.110)	(0.090)	(0.666)	(0.455)	(0.371)
40 yrs, married with children	0.003	-0.340	-0.012	-0.013	-0.143	-0.018*	0.021	-0.358	-0.010
	(0.764)	(0.427)	(0.225)	(0.275)	(0.819)	(0.087)	(0.165)	(0.540)	(0.589)
55 yrs, single, no children	0.016*	0.652	0.006	0.008	-0.051	-0.002	0.020	0.853	0.015
	(0.069)	(0.109)	(0.535)	(0.471)	(0.934)	(0.826)	(0.142)	(0.108)	(0.394)
55 yrs, married, no children	-0.028***	-0.966***	-0.018**	-0.017**	-0.809*	-0.013	-0.035***	-0.950**	-0.013
	(0.000)	(0.002)	(0.032)	(0.042)	(0.093)	(0.167)	(0.001)	(0.020)	(0.384)

p-values in parentheses; testing the null whether linear combination of coefficients equal to zero. * significant at 10%; ** significant at 5%; *** significant at 1%. (1) Employment dummy as dependent variable, FE; (2) Working hours per week as dependent variable, FE IV; (3) Log working hours per week as dependent variable, FE IV. In (2) and (3) wages are also included in the regression; in these cases effects are calculated at the value of the average wage in the sample (all, men or women).

Summing up, several interesting results emerge when we interact personal characteristics with windfall gains. First, the effects of windfall gains operate both at the external and internal margin, but they tend to be stronger at the external margin. This suggests that after receiving unanticipated windfall gains, people adjust their labour supply mainly by dropping out of the labour force, rather than by reducing their hours worked. Second, there is evidence that for some individuals (e.g., for single middle-aged individuals), effects on working hours can be positive. And third, for young and old people as well as individuals who are married with young children, windfall gains tend to have the most negative impact on labour supply.

6 Conclusion

In this chapter, we investigate whether European workers adjust labour supply (that is, labour market participation and working hours) in response to windfall gains. According to the life-cycle model of labour supply, unanticipated gains in non-earned income are expected to have negative effects on labour supply (MaCurdy, 1981). We use information from the European Household Panel for the 1994 to 2001 period to shed some light on the question of interest.

We find weak evidence that individuals react to windfall gains by reducing their working hours. The effects, however, seem to be small, which is in line with some previous findings (Holtz-Eakin, 1993; Joulfaian and Wilhelm, 1994; Henley, 2004). There is evidence that individuals adjust their labour supply mostly along the external margin, by dropping out of the labour force, rather than by reducing their working hours while staying in employment. Furthermore, we report that the effects on labour supply are stronger in the case of large windfall gains. The results and conclusions are robust to different specifications and they are not affected by the choice of which measure of wage is used on the right hand side: the hourly wage or the net monthly wage.

Finally, when we allow for heterogeneity of the effects across individuals, the results suggest that the impact of windfall gains on labour supply: (i) is more important for young and old individuals, (ii) is most negative for married individuals with young children, (iii) but can be positive for single individuals aged around 40 years. The last effect can be explained by the effect of windfall gains on reducing liquidity constraints in capital markets and thus encouraging people to become self-employed, as suggested in Lindh and Ohlsson (1996) and Taylor (2001) and confirmed also in our data.

7 References

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Appendix

Appendix Table 1: Effects of windfall gains on probability of becoming self-employed

Dependent variable: Self-employed dummy						
	all	men	women	all	men	women
	(1)	(2)	(3)	(4)	(5)	(6)
windfall gains dummy	0.00807***	0.0122***	0.00399			
	(0.00231)	(0.00358)	(0.00297)			
small windfall gains dummy				0.00198	0.00599	-0.00195
				(0.00294)	(0.00454)	(0.00378)
medium windfall gains dummy				0.00900***	0.0147***	0.00345
				(0.00317)	(0.00492)	(0.00406)
large windfall gains dummy				0.0144***	0.0175**	0.0120*
				(0.00535)	(0.00832)	(0.00685)
age	0.0208***	0.0312***	0.0114***	0.0212***	0.0316***	0.0117***
	(0.00175)	(0.00272)	(0.00222)	(0.00174)	(0.00272)	(0.00222)
age^2	-0.000243***	-0.000359***	-0.000139***	-0.000245***	-0.000361***	-0.000140***
	(1.99e-05)	(3.09e-05)	(2.53e-05)	(1.99e-05)	(3.09e-05)	(2.53e-05)
married dummy	0.0219***	0.0335***	0.00945*	0.0219***	0.0334***	0.00944*
	(0.00437)	(0.00680)	(0.00560)	(0.00437)	(0.00680)	(0.00560)
children 0-6 in household dummy	-0.000748	-0.00464	0.00135	-0.000758	-0.00475	0.00147
	(0.00367)	(0.00552)	(0.00486)	(0.00367)	(0.00552)	(0.00487)
children 7-15 in household dummy	-0.00284	-0.00471	-0.00221	-0.00280	-0.00469	-0.00214
	(0.00333)	(0.00516)	(0.00429)	(0.00333)	(0.00516)	(0.00429)
Constant	-0.316***	-0.490***	-0.156***	-0.328***	-0.502***	-0.169***
	(0.0378)	(0.0591)	(0.0482)	(0.0377)	(0.0589)	(0.0481)
Observations	54210	26177	28033	54210	26177	28033
Number of individuals	10397	5088	5309	10397	5088	5309
R-squared	0.005	0.011	0.002	0.006	0.011	0.002

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include individual fixed effects.

Appendix Table 2: Effects of windfall gains on probability of becoming self-employed (with interactions)

Dependent variable: Self-employed dummy			
	all	men	women
	(1)	(2)	(3)
windfall dummy	-0.0731*	-0.0709	-0.0698
	(0.0374)	(0.0580)	(0.0483)
age	0.0159***	0.0283***	0.00510
	(0.00243)	(0.00379)	(0.00311)
interaction of windfall gains with age	0.00465**	0.00449	0.00435*
	(0.00187)	(0.00290)	(0.00241)
age^2	-0.000182***	-0.000322***	-6.07e-05
	(2.91e-05)	(4.52e-05)	(3.73e-05)
interaction of windfall gains with age^2	-6.13e-05***	-5.84e-05*	-5.82e-05**
	(2.22e-05)	(3.44e-05)	(2.87e-05)
married dummy	0.0201***	0.0257***	0.0114*
	(0.00508)	(0.00801)	(0.00645)
interaction of windfall gains with married dummy	0.00361	0.0139*	-0.00218
	(0.00447)	(0.00730)	(0.00549)
children 0-6 in hhold dummy	0.00358	0.00496	0.000725
	(0.00476)	(0.00724)	(0.00625)
interaction of windfall gains with children 0-6 dummy	-0.00653	-0.0161**	0.00311
	(0.00503)	(0.00770)	(0.00661)
children 7-15 in hhold dummy	0.00251	0.00485	-0.00122
	(0.00435)	(0.00678)	(0.00558)
interaction of windfall gains with children 7-15 dummy	-0.00808*	-0.0149**	-0.000610
	(0.00486)	(0.00758)	(0.00625)
Constant	-0.225***	-0.436***	-0.0395
	(0.0492)	(0.0768)	(0.0629)
Observations	54210	26177	28033
Number of individuals	10397	5088	5309
R-squared	0.006	0.011	0.002

Standard errors in parentheses, * significant at 10%; ** significant at 5%; *** significant at 1%. All regressions include individual fixed effects.

General Conclusion

This work consists of three chapters on the general topic of working hours. I study working hours in relation to childcare support, wage inequality and windfall gains.

In Chapter I test whether fiscal family-support policies can help explain the differences in working-hours across countries. This has been suggested in the literature by Rogerson (2007) and Ragan (2005), but has so far relied on simulations of stylised macroeconomic models with a representative agent. However, the proposed explanations imply some important effects at a more disaggregated level. I therefore focus on differences between people with children and people without children.

I consider two types of public family policies: childcare subsidies and family cash benefits. With a simple theoretical model I show that we would expect the effect of childcare subsidies on aggregate working hours 1) to be positive in general, 2) to be positive for both parents and nonparents, but 3) be stronger for parents. There should be 4) no aggregate effects of family cash benefits on working hours, 5) a negative effect for parents, and 6) a positive effect for nonparents. The effects of family policies are also expected to affect the time in childcare within a household. According to the model outlined in this chapter there is 1) a negative effect of childcare subsidies on aggregate time spent in childcare in general, 2) negative for both parents and nonparents, but 3) stronger for parents. There should be 4) positive aggregate effects of family cash benefits on childcare time, 5) the effects should be positive for parents, and 6) negative for nonparents.

I test these predictions using cross country variation from European Household Panel and US CPS data. In the aggregate setting there is no supportive evidence for the idea that family policies help explain differences in working hours across countries. In preferred specifications with other controls included, the effects are close to zero and insignificant. In countries with higher childcare subsidies, participation in the labour force is indeed higher, but working hours conditional on working are lower, bringing the aggregate effect of policies to zero. When comparing the effects of parents and nonparents, the differences in the coefficients are counterintuitive. There is an indication that results go in the right direction for females, but are cancelled out by the counter effects on males; hence, on the whole, family policies do not have the expected effect on working hours.

Furthermore, in regressions with time spent in childcare at home on the left hand side, the effects of policy variables contradict the theory. Childcare subsidies, for example, seem to increase the time in childcare at home and this effect seems to be, counter-intuitively, strongest for parents.

An analysis performed on the sample of people older than 55 years, for which family policy should not matter much, seems to indicate that childcare subsidies tend to reduce the time in childcare for this group. One possible explanation for this is that public childcare support actually relieves grandparents from childcare, rather than parents. But, because older people do not participate in the labour market very much, family support cannot have a strong impact on working hours. Finally, I incorporate into the analysis two measures of cultural factors: the average number of adults in the household and the incidence of part-time work in the country. After controlling for these two variables, the aggregate effects of childcare subsidies on working hours become positive and significant; however, the effects on parents are still weaker than the effects on nonparents and the results from the analysis on the time in childcare spent at home still contradict the theory.

The lack of empirical evidence in support of the idea that family policies can help explain the differences in working hours across countries is striking. The family policy story can perhaps be used narrowly to explain the differences between Sweden and the US, but it does not bear the inclusion of a greater set of countries and a greater set of controls. Nor does it bear the separation of effects for parents and nonparents. I therefore conclude that the family policy story contributes little to the explanation of the differences in working hours across countries.

However, it is important to note that in practice, family policies are not implemented solely to affect labour supply. Instead, they may be directed at other objectives such as gender equity and fertility, for example, and some of these can affect labour supply in their own way, possibly masking the effects of child subsidies and cash benefits. Due to data limitations they have not been included in the analysis above, but they should be addressed in future research.

The objective of this chapter is to test whether the predictions of macroeconomists trying to explain variation in working hours across countries are correct, hence the model used is a macro type model and includes only minor variation across individuals regarding preferences and public policies. It is possible, however, that very simple assumptions in such macro models give rise to effects that would not appear in more involved models. It has been shown that in models that take into account labour supply, fertility and childcare decisions, the effects of childcare policies are mostly of an ambiguous sign. Therefore, one could conclude that in this particular instance, the hypotheses of macroeconomists are not grounded in microeconomic behaviour, a conclusion that is in line with the empirical conclusions presented in this chapter.

Finally, it is important to stress that given the slow variation in policies over time, the analysis is done mainly by exploiting variation across countries. However, the data is available only for a relatively small number of countries. With more variation in public policies over time, or when data becomes available for a greater number of countries, future analysis may lead to more robust conclusions.

In Chapter II I argue that a mean-preserving spread in offered wages (rising wage inequality) lowers average working hours in the economy. The exact mechanism works through the concavity of the labour supply. Intuitively, with wage inequality rising, the wages of highly paid workers rise and wages of low paid workers fall. Due to the concavity in the labour supply, responses to the changes in wages are stronger at the bottom of the wage distribution, hence the working hours of low-paid workers decrease more than working hours of high-paid workers rise. Some of the low-paid workers may also decide to drop out of the labour force and become inactive, thus reducing their labour supply to zero. Therefore, as a consequence, average working hours in the economy fall.

This explanation implicitly assumes that working hours are just a response to labour market opportunities. In this I follow the labour supply literature such as Juhn et al. (1991), Juhn (1992) and Welch (1997). The contribution of Chapter II is that it explicitly recognizes that changes in the inequality of offered wages have effects on average working hours in the economy. Using the CPS-MORG data for prime-age men I find evidence for this explanation. First I establish empirically the concavity of the labour supply function. Secondly, using various variables to measure wages and labour supply, and using several different specifications, I find empirical evidence that after controlling for the average wage, wage inequality has indeed a negative and significant effect on the labour supply. According to my results, the increase in wage inequality that occurred in the US over the 1979-2008 period would cause a 3% decrease in average hours worked in the economy. Rising inequality can thus potentially explain 30-40% of the decrease in working hours over time.

The results in Chapter II are interesting for students of labour supply, for students of inequality and for researchers interested in explaining differences in working hours across countries and over time. Moreover, the results are also interesting from the point of view of public policies that affect wage inequality.

In Chapter III my co-author and I investigate whether European workers adjust labour supply in response to windfall gains. According to the life-cycle model of the labour supply, unanticipated gains in non-earned income are expected to have negative effects on labour supply (MaCurdy, 1981). We use information from the European Household Panel for the 1994 to 2001 period to shed some light on the question of interest.

We find evidence that individuals react to windfall gains by reducing their working hours. The effects, however seem to be small, which is in line with some previous findings (Holtz-Eakin, 1993; Joulfaian and Wilhelm, 1994; Henley, 2004). There is evidence that individuals adjust their labour supply mostly along the external margin, by dropping out of the labour force, rather than by reducing their working hours while staying in employment. Furthermore, we report that the effects on labour supply are stronger in the case of large windfall gains.

Finally, when we allow for heterogeneity of the effects across individuals, the results suggest that the impact of windfall gains on labour supply: (i) is more important for young and old individuals, (ii) is most negative for married individuals with young children, (iii) but can be positive for single individuals at the age of around 40 years. The last effect can be explained by the effect of windfall gains on reducing liquidity constraints in capital markets and thus encouraging people to become self-employed, as suggested in Lindh and Ohlsson (1996) and Taylor (2001) and confirmed also in our data.

At the end I will give a brief critical assessment of the work that has been done and discuss some potential avenues for future research.

In Chapter I, I initially expected, consistent with a large body of macroeconomic theory, that family policies would turn out to have an effect on aggregate labour supply. However, it turns out that I am left with proving that cash benefits to families and childcare subsidies have no impact on aggregate labour supply across a group of countries. As an empirical question, it is an uphill battle to prove that some factor has zero effect, as any test of this sort essentially has no power, therefore such analysis is very much prone to criticism. One must be confident that the model is appropriately specified, that the measures of the policy variables are good and that one has controlled adequately for other policies and cultural differences that might affect employment and working hours decisions. Furthermore, it is difficult to tackle a macroeconomic question with individual data, because even though the large number of individuals yields many data points, the country level variables still suffer from a limited number of degrees of freedom.

Nevertheless, the question being asked in Chapter I remains an interesting one and more research can be done on it. The most obvious thing to do is to wait for a longer data series, which would hopefully bring more variation into the policy variables. Similarly, with data available for a greater set of countries, the degrees of freedom problem will become less severe. However, some research could be done also with existing data. For example, instead of individuals, analysis could be done on households. Empirical work on working hours could also rely on testing more complex models of household behaviour, following household behaviour regarding labour supply, fertility and childcare more closely.

In Chapter II I examine changes in working hours in the US for the 1979-2008 period using data on prime-age men. I exclude women due to strong growth in female participation during this period and due to complications of female labour supply associated with childbearing. But, trends in the labour supply of men and women might be related, resulting in reductions of male labour supply precisely due to increases in female labour supply. This could occur if women are substitutes for men at low wages, or if husbands can remain inactive because wives are now earning a salary. Hence, further research should include female labour supply and delve into exploring the interrelations between trends in male and female labour supply. One obvious line of research would be to test the predictions of Chapter II, that inequality reduces average hours of work, with data from a different country. Another interesting area of research would also be to explore the complementarities (or substitutions) of labour supply within households with changing labour market opportunities.

In Chapter III I test the effects of receiving unexpected windfall gains on labour supply, relying on predictions from a life cycle model. Again my unit of observation is the individual, but I could as well explore the effects on the labour supply for households as a whole and potential changes and substitutions within a household. It would be very interesting to see the effects on other uses of time, such as leisure, childcare, and on consumption patterns. The variable measuring windfall gains in my data is a household level variable and is available in three discreet intervals, only indirectly indicating size of the gain. Ideally, however, one should find data where it is clear who the receiver of the windfall gain is and what the amount of the gain is.