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Does Worksharing Work? Some Empirical Evidence from the IAB Panel

M.J. ANDREWS, T. SCHANK UND R. SIMMONS

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Does Worksharing Work?
Some Empirical Evidence from the IAB Panel

Martyn J. Andrews[a], Thorsten Schank[b], Robert Simmons[c]

Abstract: Recent policy debate in Europe suggests that a shorter workweek will lead to more jobs (worksharing). We derive and estimate a model where the firm employs two types of worker, some working overtime, the rest standard hours. Worksharing is not always a prediction of the theory. Using German establishment-level panel data (the IAB panel), 1993–1999, we find pro-worksharing effects in small plants in the East German non-service sector. There is evidence that a cut in standard hours lowers the proportion of overtime workers in a plant, as predicted by the theory, and increases the proportion of standard-time plants.

Zusammenfassung: In der wirtschaftspolitischen Debatte werden immer wieder Verkürzungen der wöchentlichen Normalarbeitszeit zur Bekämpfung der Arbeitslosigkeit gefordert. Wir präsentieren in dem vorliegenden Papier ein Arbeitsnachfragemodell, das zwischen Beschäftigten mit und ohne Überstunden unterscheidet. Der Effekt einer Arbeitszeitverkürzung auf die Beschäftigung kann dabei nicht eindeutig beantwortet werden. Auf der Basis des IAB-Betriebspanels, 1993–1999, finden wir, dass in kleinen Betrieben des Produzierenden Gewerbes Ostdeutschlands Arbeitszeitverkürzungen und Beschäftigungserhöhungen miteinander einhergingen. Außerdem zeigt sich, dass eine Verkürzung der wöchentlichen Arbeitszeit den Anteil der Überstundenbeschäftigten – so wie von der Theorie vorhergesagt – senkt und den Anteil der Betriebe ohne Überstunden erhöht.

Keywords: worksharing; plant-level panel data; Germany

New JEL-Classification: C23, C24, J23

[a]Dr. Martyn J. Andrews, School of Economic Studies, University of Manchester, Manchester, M13 9PL, United Kingdom, martyn.andrews@man.ac.uk.

[b]Dr. Thorsten Schank, Friedrich-Alexander-Universität Erlangen-Nürnberg, Lehrstuhl für Arbeitsmarkt- und Regionalpolitik, Lange Gasse 20, D-90403 Nürnberg, thorsten.schank@wiso.uni-erlangen.de.

[c]Dr. Robert Simmons, The Management School, Department of Economics, Lancaster University, Lancaster, LA1 4YX, United Kingdom, r.simmons@lancaster.ac.uk.

1 Introduction

One of the more contentious issues concerning labour-market policy is whether a sustained reduction in the length of the working week leads to more jobs. This policy is known as worksharing, because a fixed number of hours worked in the economy are spread over more workers. It is contentious because there are widely differing views promulgated by politicians, trade unions and economists as to whether it actually works. It is in continental Europe where the policy is debated most; two examples of worksharing programmes that illustrate the scope for disagreement are in France and Germany.[1] In France, a complex policy programme involving reductions in the working week (hereafter, 'workweek' or 'standard hours') was initiated under the Aubry laws of 1998 and 2000. These laws provided for a statutory standard workweek of 35 hours down from 39, initially for workers in organisations employing 20 or more and then extended to cover smaller workplaces. Firms were encouraged to adopt a shorter basic workweek, with reduced social security contributions as an incentive, and to place limits of 130 hours per year per worker on the amount of overtime performed. French survey evidence does indeed show a substantial cut in average weekly hours worked since 2000 when the policy was first applied. According to Government figures, during 2000 and 2001, 83,000 companies moved to a 35 hour workweek and bargaining agreements involving cuts in the workweek are credited with saving or creating a total of 365,000 jobs over this period (European Industrial Relations Review 2001). Also, a Government survey conducted over the period November 2000 to January 2001 showed that the majority of workers felt their quality of life had improved as a result of the introduction of the 35 hour week (European Industrial Relations Review 2001). However, the move to a shorter workweek was attacked by some Government ministers and employers' federations as contributing to higher unit labour costs, lower French competitiveness and weak economic growth.[2] As part of an employer-led backlash against the Aubry laws, the overtime limit was raised to 180 hours and application of the laws made less stringent.[3]

In Germany, which is the focus for this paper, the dominant union, IG Metall, secured a series of collective bargaining agreements though the 1980s and 1990s to take the

[1]Elsewhere, the Italian government announced a similar move in October 1997, although the legislation has not been introduced yet. In 1999, unions in Greece, Spain, Sweden, Belgium and Switzerland called for a similar policy, although in no cases have statutory economy-wide cuts in hours been implemented; rather any cuts in hours have come through bargaining agreements. Other measures to induce worksharing via a reduction in hours worked are to increase the yearly holiday entitlement, to limit the maximum permitted overtime hours or to increase the overtime premia. Increasing the overtime premium is the only provision in the United States to induce worksharing, where the Fair Labor Standards Act mandates an overtime premium of 50 percent to be paid after 40 hours.

[2]Jean-Paul Fitoussi is quoted as saying that it is "not possible for eight people working for six hours to produce the same amount as six people working eight hours" (The Guardian 2003).

[3]See European Industrial Relations Review (2002*a*, 2002*b*).

workweek to 35 hours in 1995, which have been followed in other industries.[4] In June 2003, though, an attempt by IG Metall to obtain a 35 hour workweek in East German metalworking plants, in order to gain parity with West Germany, failed as employers' organisations mounted a successful campaign of resistance to strike action called by the union. The East German employers argued that parity of the workweek with the West would undermine a useful cost advantage for employers in the East and cited potential job losses as an important part of their argument.[5]

Clearly, policy makers still consider worksharing a tool for reducing unemployment in Europe. This is despite the fact that many economists have argued that worksharing does not work. The basic idea is that, as the workweek is cut, firms substitute hours by workers, with workers receiving lower weekly pay (income-sharing) but not being worse off providing they value leisure sufficiently. However, there is a well-known problem with the theory (Hart 1987, Calmfors & Hoel 1988, Leslie 1991, Hamermesh 1993, Hunt 1999). The policy only works for firms who do not offer overtime. For those that employ exclusively overtime workers, cutting standard hours makes hours worked per person more expensive, and both scale and substitution effects predict that the firm demands *fewer* workers. Not only is the theory *a priori* ambiguous, because it depends on whether firms employ overtime workers, it illustrates clearly the fallacy of fixed lump of labour: firms will demand fewer man-hours if labour costs go up.

The empirical evidence is equally unconvincing. However it based almost exclusively on aggregate- or industry-level data; as yet there is virtually no demand-side micro-econometric evidence with which to assess this issue (see Hart & Wilson (1988), Hunt (1999), Hernanz, Izquierdo & Jimeno (1999) and Crépon & Kramarz (2002) and Section 2 below). In this paper, we present empirical evidence using a panel of German plants (the IAB panel) for the period 1993–99. We estimate various employment regressions to see whether standard hours have a negative impact.

There are two features of our data that ensure our empirical results add value to the existing literature. First, not only do we have plant-level data, they are also a panel, which means that we can control for unobserved scale effects which contaminate many studies of labour demand. Second, uniquely, we observe the proportion of overtime workers in a given plant. Given the above problem with the theory, it is essential to distinguish between plants who offer overtime and those who do not, that is examine whether estimated employment-standard hours elasticities vary with—in fact, change sign with—the plant's working-time regime. In reality, many plants employ both overtime and non-overtime workers, which means we can also estimate the effect of standard hours on this proportion, thereby decomposing employment-standard hours elasticities into substitution and scale effects.

[4] See Hunt (1998) for an excellent survey. A later demand for 32 hours was unsuccessful.

[5] For a full review of this case, see European Industrial Relations Review (2003*a*, 2003*b*).

Because the theoretical literature only analyses firms who employ exclusively standard-time workers or exclusively standard-time workers, we inform our empirical analysis by developing a model first suggested by Leslie (1991) which incorporates both types of worker within the same firm. In this so-called compositional model, the theory predicts that, as standard hours are cut, the proportion of overtime workers in the workforce falls as firms substitute standard-time workers for overtime workers. The model incorporates the variables we have at our disposal in the data.

The paper is organised as follows. In Section 2, we survey the existing empirical evidence as to whether cuts in the workweek have indeed led to, or are associated with, increases in employment and/or actual hours of work. In Section 3, we extend standard models of labour demand by modelling the firm's choice of overtime regime endogenously and then present our own compositional model that incorporates standard time and overtime workers in the same firm. In Section 4 we describe the data. In Sections 5 and 6 we discuss some econometric issues and present our regression results from the IAB panel. Section 7 concludes.

2 Evidence

In this section we survey the existing empirical evidence as to whether cuts in standard hours, $\bar{H}$, have led to, or are associated with, increases in employment, N, hours of work, H (or, equivalently, overtime, V), and the proportion of individuals working overtime. Often these estimates are interpreted as the effect of standard hours on the firms factor demands for hours and employment, but reduced-form interpretations are also admissible.

Almost all the early evidence came from time-series data; see Schank (2001, Table 3.1) for a comprehensive summary of 17 studies covering 6 industrialised economies. For most countries, time-series plots of H and $\bar{H}$ indicate secular declines in both series, hence the strong positive correlation in time-series studies. If overtime is solely a short-run, disequilibrium phenomenon (i.e. $\eta_{V\bar{H}} \approx 0$), this implies that a unit elasticity is an upper limit for long-run $\eta_{H\bar{H}}$.[6] In fact, most estimates are close to unity; exceptions are Brunello (1989) for Japan and König & Pohlmeier (1989) for West Germany. In time-series studies where employment regressions are also run, estimates of$\eta_{N\bar{H}}$ are not so well-determined. A negative relationship between employment and standard hours is found in European economies, e.g. Finland, Italy, Netherlands and West Germany. Again Japan is an exception. More recently, Kapteyn, Kalwij & Zaidi (2000) estimate a six-variable VAR on a panel for 13 OECD countries, and find no causal relationship

[6] Throughout this paper the elasticity of y with respect to x is denoted η_{yx}. If it is unambiguously negative then it might written as a positive quantity: $|\eta_{yx}|$.

from actual hours worked to employment. Jacobson & Ohlsson (2000) do the same for Sweden, with the same finding.

In the late 1990s, the need for micro-econometric evidence, using either firm/plant- or individual-level panel datasets, became paramount, there being only Hart & Wilson (1988) hitherto. Table 1 reports what evidence there is, almost exclusively for the UK or Germany. Hart & Wilson (1988) estimate conditional factor demand schedules for hours of work and employment from a panel of 52 engineering enterprises for 5 years (1978–1982) in the UK, and find that $\eta_{N\bar{H}}$ is badly determined, because it conflates different effects for 9 firms who offer no overtime with the rest who do. When the sample is stratified by working-time regime, the elasticity is -0.49 for the zero overtime case but is 0.41 for firms who offer overtime. This supports the prediction that the two types of firm respond differently to changes in $\bar{H}$.

Only three other studies have been conducted. First, Hunt (1999) constructed an 30 industry-level bi-annual panel from the West German Mikrozensus for 1982–93, and so was able to control for industry fixed effects, and found positive but insignificant effects. The second is Hernanz et al. (1999), whose preliminary estimates imply an insignificant effect of standard hours on employment for Spanish manufacturing firms, irrespective of specification and estimation method (for removing plant-level fixed effects and instrumenting standard hours), thereby confirming the previous two studies.

The last, most compelling, evidence comes from France. Crépon & Kramarz (2002) examine the impact on transitions from employment to non-employment of the introduction of unanticipated legislation in 1982, which reduced the workweek from 40 to 39 hours. Using the French Labour Force Survey, a panel that records individual labour-market transitions, Crépon & Kramarz find very strong evidence against work-sharing, where the probability of leaving employment increases by between 2% and 4%, depending on specification. The effects are even bigger for individuals on minimum wages, where firms are unable to offset increased overtime costs by wage cuts, which were ruled out by law.

All the micro-econometric evidence for $\eta_{H\bar{H}}$ confirms the robust result found in time-series data. This is robust to whether censored, truncated, fixed-effects or OLS regressions are run. Also note that corresponding probit estimates consistently suggest a positive effect of a cut in standard hours on the probability of working overtime (an estimate of -0.005 converts to a semi-elasticity of -0.20, multiplying by a standard 40-hour week).

Taking all the evidence on $\partial H/\partial\bar{H}$ together, we conjecture that it is sufficiently convincing to labelled a Stylised Fact:

The effect of standard hours on actual hours is positive, with an elasticity robustly estimated at slightly less than unity.

However, as noted, the effect of standard hours on employment has yet to be established empirically, especially using firm- or individual-level panel data. Our own empirical work below uses a German panel of plants, where the proportion of overtime workers in a plant is information not analysed in the literature hitherto.

3 The theory of the firm

In the Introduction it was stated that worksharing—a cut in the workweek leading to a more employment—is not a prediction of the standard theory of the firm, but may occur only if the firm optimally chooses hours of work at the exogenously given workweek. In this Section we extend standard models of labour demand by modelling the firm's choice of overtime regime endogenously (Sections 3.1 and 3.2). This is done by amending standard labour supply techniques for maximising utility with piecewise-linear constraints (see, for example, Blundell & MaCurdy 1999, Section 6.4.1). Because firms do not employ exclusively overtime workers or standard time workers, we develop our own compositional model in Section 3.3.

Consider a firm free to choose both the level of employment, N, and weekly hours, H, per employee. All workers work the same number of hours. The firm's cost function is given by

$$C = N(wH + z) \qquad \text{if } H \leq \bar{H}; \tag{1}$$

$$C = N[w\bar{H} + \gamma w(H - \bar{H}) + z] \qquad \text{if } H \geq \bar{H}. \tag{2}$$

Weekly hours may be greater than the standard workweek, in which case overtime hours $V \equiv H - \bar{H}$ are strictly positive. Each hour up to $\bar{H}$ is paid w; overtime hours are paid at premium rate γw, where $\gamma > 1$. z represents quasi-fixed labour costs, i.e. those fringe costs which are independent of hours worked, imputed on a per-period basis. (Typically they represent hiring and firing costs.)[7] The isocost contour in (N, H) space comprises two convex segments which form a kink at $H = \bar{H}$ (e.g. $A_0B_0D_0$ in Figure 1). The firm's (strictly concave) revenue function is denoted by $\theta R(H, N)$, where θ is a demand shock.

Solutions are found in two steps. First, we determine the choice of H and N conditional on locating on a particular segment or kink. The profit-maximising firm faces different

[7] This piece-wise linear in hours cost schedule is the simplest way of incorporating overtime and is a reasonable approximation to reality. For the UK, Bell & Hart (1999) report that γ is invariant to the number of overtime hours in the 1996 NES and find that $\gamma = 1.4$.

costs for both segments. More precisely, if $H \leq \bar{H}$, the firm chooses H and N to maximise $\theta R(H, N) - C$, where C is given in Equation (1); if $H \geq \bar{H}$, the firm solves the same problem, except that costs are now given in Equation (2). In general, these solutions are written:

$$\left.\begin{array}{l} N = N_0(w/\theta, z/\theta) \\ H = H_0(w/\theta, z/\theta) \end{array}\right\} \quad \text{if } H \leq \bar{H}, \text{ i.e. the undertime segment;} \tag{3}$$

$$\left.\begin{array}{l} N = N_1(w/\theta, \bar{H}, z/\theta) \\ H = \bar{H} \end{array}\right\} \quad \text{if } H = \bar{H}, \text{ i.e. the kink;} \tag{4}$$

$$\left.\begin{array}{l} N = N_2(w/\theta, \bar{H}, z/\theta) \\ H = H_2(w/\theta, \bar{H}, z/\theta) \end{array}\right\} \quad \text{if } H \geq \bar{H}, \text{ i.e. the overtime segment.} \tag{5}$$

The second step is to determine the segments or kinks on which the firm operates. If

$$H_0(w/\theta, z/\theta) < \bar{H},$$

then the firm operates on the lower segment, i.e. offers all its employees hours of work lower then the standard workweek. We ignore this solution henceforth, because short-time working is not observed very often (part-time working is usually seen as a supply-side phenomenon). Next, if

$$H_2(w/\theta, \bar{H}, z/\theta) < \bar{H} \leq H_0(w/\theta, z/\theta),$$

then the firm operates on the kink. All its employees work the standard workweek. Firms whose working-time arrangements are characterised as such are hereafter labelled as 'standard-time firms'. Finally, if

$$H_2(w/\theta, \bar{H}, z/\theta) \geq \bar{H}, \tag{6}$$

then the firm operates on the upper segment and all its employees work overtime. These firms are hereafter labelled 'overtime firms'. The two possible solutions (or 'working-time regimes') are drawn in Figure 1. For $A_0B_0D_0$, these are points B_0 (standard-time firms) and C_0 (overtime firms) respectively.

Notice that not only do the arguments of these functions vary between the three sets of solutions, so do the functions themselves. In particular, the effect of the workweek on employment varies between overtime and standard-time firms. We now discuss this important issue more fully.

3.1 Overtime firms

The analysis of changes in the workweek where all firms offer overtime is well established (Hart 1987, Leslie 1991, Hamermesh 1993). The demand for employment and hours functions are given by Equation (5), whose properties depend, in part, upon the underlying technology generating the revenue function $\theta R(H, N)$. Here, we follow the strong tradition in the literature which suggests that hours worked are independent of scale (Ehrenberg 1971).[8] With this assumption, a revenue shock does not alter the profit-maximising demand for hours but does affect the demand for employment positively. The 'no scale effects' revenue function is $\theta R(g(H)N)$ where $g()$ is any concave function. It is easily shown that the comparative statics properties of the two demand functions are that $N_{w/\theta}$ is ambiguous in sign, and that $N_{\bar{H}} > 0$, $N_{z/\theta} < 0$, $N_\theta > 0$, $H_{\bar{H}} < 0$, $H_{w/z} < 0$ and $H_\theta = 0$. For a firm to optimally offer overtime to all its employees, Equation (6) suggests that it must face relatively low standard hours. In Appendix A.1 we show that, for a Cobb-Douglas production function, the appropriate condition states that the share of fixed costs in 'normal' costs $k \equiv z/(w\bar{H} + z)$ must be sufficiently high (and, indeed, implausibly so).

A cut in the workweek, $\bar{H}$, is qualitatively the same as an increase in fixed costs per worker, z. For given output, the marginal cost of an employee (the so-called extensive margin) rises but the marginal cost of an overtime hour (the intensive margin) remains constant, and so the firm substitutes away from employment towards hours (C_0 to C_1 in Figure 1). Effectively, $-\bar{H}$ acts as the own price of employment. With profit maximisation, there is an additional scale effect, whereby the firm demands fewer hours and employees, because costs have risen. For the 'no scale effects' revenue function, the solution lies directly below C_1. The overall effect of a cut in the workweek on employment is unambiguously negative ($N_{\bar{H}} > 0$), being negative for both income and substitution effects: work-sharing is fundamentally flawed. There is more expensive overtime as the workweek is cut. Crépon & Kramarz (2002) is an overtime model of this type, where, recall, they find strong evidence against worksharing.

Clearly a theoretical model where everyone works overtime is not what proponents of worksharing have in mind. Moreover, this model is completely at odds with the empirical evidence, where the effect of a cut in the workweek is to reduce actual hours almost hour for hour (see the Stylised Fact in Section 2 above). The theory requires that $\partial V/\partial \bar{H} < -1$, whereas the estimated elasticities are zero or slightly negative.[9]

[8] To quote Hamermesh (1993) "After all, there is no evidence that the weekly hours of full-time workers at General Workers differ substantially from those at the local steel fabricator".

[9] There have been many attempts at resolving this conflict between theory and evidence. Some authors have focused on imposing a particular structure on the production technology (Hart 1987, Calmfors & Hoel 1988, Leslie 1991). Others specify the overtime premium γ as a increasing function of hours worked; Hart subsequently found that, in the data, γ increases with hours in Japan, but not in the UK or US (Hart & Ruffell 1993, Hart, Malley & Ruffell 1996, Bell & Hart 1999).

3.2 Standard-time firms

If it is optimal for the firm to operate at the kink, effectively employment is chosen conditional on the exogenously determined workweek, $H = \bar{H}$.[10] The firm's problem can be more simply stated as

$$\max_{N} \theta R(\bar{H}, N) - (w\bar{H} + z)N. \tag{7}$$

This generates the labour demand Equation (4). The variables that enter are the same as for the overtime regime (see Equation 5); it is the comparative static effects that are different: $N_{w/\theta} < 0$, $N_{\bar{H}} \lesseqgtr 0$, $N_{z/\theta} < 0$, and $N_\theta > 0$. It is clear from (7) that $\bar{H}$ is a price of employment, in direct contrast to the overtime model above. Just like a cut in the wage, a cut in $\bar{H}$ *lowers* the marginal cost of an extra employee. On its own this increases employee demand. However a cut in $\bar{H}$ also affects marginal revenue; only if the cross-partial is sufficiently positive, ie marginal revenue falls by more than marginal cost, does employment actually fall. In the Cobb-Douglas case (see Appendix A.2) all that is needed for worksharing to work is that $0 < k < 1 - \alpha$, where α is the exponent on H.

Figure 1 illustrates the pure substitution effect of such a cut in the workweek, from B_0 to B_1, and makes it quite clear why the effect is opposite to the interior solution. Also notice that because costs have risen, the scale effect moves the solution below B_1 (vertically if there are no scale effects), which illustrates the ambiguity in the partial derivative.

It is this model that the proponents of worksharing had in mind: the firm has to employ more workers if hours are exogenously cut. As already noted, this is not a prediction for overtime plants. Whether worksharing works depends on the extent to which firms employ overtime workers. Many workers in Continental Europe work exactly zero overtime.[11] Of course the shape of the firm's isocost schedule tends to 'attract' observations at the kink.[12] Above, we reported that there is no convincing aggregate evidence that worksharing works, which is not surprising if some firms increase employment whilst others do exactly the opposite. Clearly, it is essential to distinguish between plants who offer overtime and those who do not, that is examine whether estimated employment-standard hours elasticities vary with—in fact, change sign—the plant's working-time regime. However, very few firms employ exclusively overtime workers.

[10]This model has also been analysed by Calmfors & Hoel (1988), and is mentioned in passing by Hart (1987) and Leslie (1991). Hunt (1999) is a special case of a model analysed by Schmidt-Sorensen (1991).

[11]From the German Socio-Economic Panel (1984–97), Bauer & Zimmermann (1999) report that only 51% of full-time male West Germans (not including Civil Servants) work some form of paid overtime. In the UK, using the New Earnings Survey, the incidence of overtime has fallen from 46% in 1979 to 32% in 1999 (Kalwij & Gregory 2000).

[12]For example, in the Cobb-Douglas example above, k needs to lie in the interval $[1-\alpha/\beta, 1-\alpha/\gamma\beta]$.

3.3 Two types of worker within a firm

In the above two subsections, we model between-firm variations in hours of work in that, given variations in exogenous parameters, some firms offer hours of work at the workweek, whereas others offer overtime. However, within each type of firm, all workers are identical. In this section, we present a model of within-firm variations in hours of work: there are two types of worker, one type works exactly the workweek whereas the other type works overtime. The firm optimally chooses the numbers of both types and the number of hours for the overtime workers. This extends Leslie (1991), the only other compositional-type model in the literature.[13]

More formally, we assume that there are $\bar{N}$ workers who work exactly $\bar{H}$ hours, who are paid an hourly wage of $\bar{w}$ and who incur quasi-fixed labour costs of $\bar{z}$, and that there are $\widehat{N}$ workers who work $\widehat{H} > \bar{H}$ hours, ie strictly positive overtime V, who are paid an hourly wage of $\gamma\widehat{w}$ and who incur quasi-fixed labour costs of $\widehat{z}$. We assume that the exogenous parameters $\bar{w}$, $\bar{z}$, $\widehat{w}$, $\widehat{z}$ are such that it is optimal for the firm to choose these two particular regimes for these two types of worker (in general each can be offered undertime/on the kink/overtime, as in Section 3 above).[14] Total employment is defined as $N \equiv \bar{N} + \widehat{N}$ and the proportion of overtime workers in total employment is defined as $p \equiv \widehat{N}/N$.

The firm's revenue depends on four factors of production. In general, all we need assume is that this function is concave. However, more tractable results can be obtained if we use an obvious generalisation of the no-scale-effects revenue function used earlier, namely

$$R = \theta R(G, L) = \theta R[f(\widehat{H})\widehat{N}, g(\bar{H})\bar{N}],$$

where the functions $R(G, L)$, $f(\widehat{H})$, and $g(\bar{H})$ are all concave. One other condition is required, that the firm must be able to choose exclusively all overtime workers or all standard-time workers if it is optimal to do so, ie $R(0, L) > 0$, $R(G, 0) > 0$, but retaining $R(0, 0) = 0$. The Cobb-Douglas production function does not have this property.

The firm's problem is written:

$$\max_{\bar{N},\widehat{N},\widehat{H}} \Pi = \theta R[f(\widehat{H})\widehat{N}, f(\bar{H})\bar{N}] - \bar{N}(\bar{w}\bar{H} + \bar{z}) - \widehat{N}[\widehat{w}\bar{H} + \gamma\widehat{w}(\widehat{H} - \bar{H}) + \widehat{z}],$$

[13]Leslie has cost-minimisation and a fixed number of total hours per week.

[14]Essentially, $\widehat{z}/\widehat{w}$ has to be sufficiently 'high' and $\bar{z}/\bar{w}$ has to be sufficiently 'low'.

with first-order conditions for $\widehat{N}$, $\bar{N}$, and $\widehat{H}$ respectively given by:

$$\theta R_{\widehat{N}} = \theta R_L f = \widehat{w}\bar{H} + \gamma\widehat{w}(\widehat{H} - \bar{H}) + \widehat{z} \tag{8}$$

$$\theta R_{\bar{N}} = \theta R_G g = \bar{w}\bar{H} + \bar{z} \tag{9}$$

$$\theta R_{\widehat{H}} = \theta R_L f' \widehat{N} = \widehat{w}\gamma. \tag{10}$$

Just as in the overtime model above, the demand for hours is solved first, from the first-order conditions for overtime hours (Equation 10) and overtime workers (Equation 8). The solution for $\widehat{H}$ is basically the same as the solution for H in Section 3.1, and therefore has exactly the same properties. In particular, $\partial\widehat{H}/\partial\bar{H} < 0$ implies that $\partial V/\partial\bar{H} < -1$.

What is of interest here is whether an exogenous cut in standard hours means that the firm increases total employment (worksharing) and whether there are clear-cut substitution effects between the two types of worker. Formal solutions are given in Appendix A.3. Intuitively, we can see that there are three distinct channels by which a cut in standard hours influences the firm's demand for overtime and standard-time workers, namely:

1. Via the marginal cost of standard-time workers, $\bar{w}\bar{H} + \bar{z}$. A cut in $\bar{H}$ induces (i) substitution towards $\bar{N}$ from $\widehat{N}$ and (ii) an increase in both $\bar{N}$ and $\widehat{N}$ from a positive scale effect. In total, $\bar{N}$ unambiguously goes up and $\widehat{N}$ also goes up if the scale effect dominates the substitution effect. This is the same pro-worksharing effect as in Subsection 3.2. The proportion of overtime workers$p$ unambiguously falls for the no-scale-effect revenue function.

2. Via the marginal cost of overtime workers, $\widehat{w}\bar{H} + \gamma\widehat{w}(\widehat{H} - \bar{H}) + \widehat{z}$. A cut in $\bar{H}$ induces (i) substitution towards $\bar{N}$ from $\widehat{N}$, as above, but now (ii) a *decrease* in both $\bar{N}$ and $\widehat{N}$ from a negative scale effect. In total, $\widehat{N}$ unambiguously goes down and $\bar{N}$ also goes down if the scale effect dominates the substitution effect. This is the same counter-worksharing effect as in Subsection 3.1. Again, the proportion of overtime workers$\,p$ unambiguously falls for the no-scale-effect revenue function.

   Thus far, these predictions are clear-cut, and general. First, worksharing is more likely to work the smaller the proportion of overtime workers in the firm. Second, the proportion of overtime workers unambiguously falls. However, these effects are ameliorated by the third channel, namely:

3. Via $\bar{H}$ in the revenue function. This is directly analogous to the effect of the capital stock as an exogenous variable in the text-book short-run labour demand model. Intuitively, one might expect the demand for both types of worker to fall,

with no substitution effects. This is true for a Cobb-Douglas revenue function—admissible providing that $0 < p < 1$—but the only unambiguous prediction from the no-scale-effect revenue function is that fewer standard-time workers are employed.

Thus we conclude: (i) ultimately it is an empirical issue as to whether worksharing works, and (ii) worksharing is primarily a substitution phenomenon. (ii) implies that we should expect to observe, in the data, a clear positive effect of standard hours on the proportion of overtime workers within a plant, and on the proportion of overtime plants in the sample. This is a different prediction to the non-compositional model in Sections 3.1 and 3.2. There, plants with a lower workweek offered all their workers overtime. As the workweek increases, there is a discrete point at which the plant offers nobody overtime, as illustrated in Figure 2. Evidence from individual-level data says that overtime incidence is decreasing in the workweek (Table 1). In the IAB data used below, we observe p, N, V and $\bar{H}$ for each plant, and so these data are ideal for examining these predictions. Our empirical results are presented in the rest of this paper.

4 The IAB panel

The dataset used in our empirical work is the *Establishment Panel Data Set* collected by the Institute of Employment and Research (IAB), Nürnberg, Germany. The data cover six years (1993, 1995–99) comprising yearly interviews with approximately 8,250 plants located in the former West Germany since 1993 and an additional 7,900 (approximately) plants located in the former East Germany since 1996.[15]

There are three potential problems with these data. The first of these arises because of the particular sampling procedure used, which causes the IAB dataset to be highly stratified. The second is that a non-negligible proportion of plants did not provide information on some of the key variables. The third is attrition, because of non-response, birth and death of plants. All are discussed in Schank (2001, chapter 4), who concludes that the impact of the latter two is more or less random for the variables we use for this study. This remains true when we reinvestigate these issues using Wooldridge (2002, Section 17.7).

To address the first problem, for each plant–year we calculate population weights by defining 320 strata for each year. These are formed out of 10 plant-size categories for 16 industries for both East and West Germany. Given the way the data were collected, we believe that, within each stratum, the plants are randomly sampled and so we ascribe

[15]Not all relevant variables were recorded in 1994.

to all plants within a given stratum the same population weight ω:

$$\omega_{js} = M_s/m_s \qquad s = 1, \ldots, 320,$$

where j indexes plants within a stratum, M_s is the number of plants in the population in stratum s (taken from the Employment Statistics Register) and m_s is the corresponding number in the IAB sample.[16] The weights vary considerably, but especially across plant-size, as summarised in Table 2. For example, included in the sample are 50% of plants in the population employing more than 5,000 employees but only 0.25% of plants employing fewer than 100 employees.

For any variable X, the weighted sample mean is given by

$$\bar{X}^w = \frac{\sum_{s=1}^{S}\sum_{j=1}^{M_s}\omega_{js}X_{js}}{\sum_{s=1}^{S}\sum_{j=1}^{M_s}\omega_{js}} = \sum_{s=1}^{S}\frac{M_s}{M}\bar{X}_s, \tag{11}$$

where $\bar{X}_s$ is the unweighted sample mean of X in stratum s, and $M \equiv \sum_s M_s$ is the number of plants in the sample. Deaton (1997, p. 67) notes that provided that the sample means for each stratum are unbiased for the corresponding population means, so is the weighted mean for the overall population mean. This is why we report both weighted and unweighted means in our table of descriptives. When X refers to employment (N), it is easier to write Equation (11) as

$$\bar{N}^w = \frac{\sum_{i=1}^{M}\omega_i N_i}{\sum_{i=1}^{M}\omega_i} \tag{12}$$

where the i index simply replaces js.

Table 3 reports the means of key variables, weighted and unweighted. Whilst we observe employment (N) for each plant-year, many plants did not supply information on other variables, with wages, overtime and investment being the worst affected. This reduced the usable dataset by about one-half. Thus, the final two columns of Table 3 refer to the unweighted regression samples, whereas the rest of the table refers to the whole dataset, that is weighted means based on all the available information.

The bottom two rows of Table 3 report the numerator and denominator of Equation (12) which are, respectively, estimates of total employment in the population and the total number of plants in the population. $\bar{N}^w$ is given in the first row of the third panel, and is 18 employees. The corresponding unweighted figures are (not shown) 517 for the West and 172 for the East. This over-sampling of large plants implies that the survey covers about 0.4% of all plants in Germany but 8% of all employees, and indicates that our empirical investigations below need to take account of these highly

[16] Throughout this subsection, for notational clarity, j refers to a plant-year.

variable population weights.

Of the variables identified in the compositional model in Subsection 3.3 above, all of $\widehat{N}$, $\bar{N}$ (and therefore p), $\bar{H}$, H, and V are observed. The actual question asked of $\bar{H}$ is: "How long is the currently agreed weekly working time for full-time workers?"[17] Standard hours $\bar{H}$ are determined by collective bargaining agreements or by individual contracts.

Like the theory, p refers mainly to paid overtime, although its definition improves slightly in 1999. In 1999, p and V refer only to those employees who work paid overtime, whereas for 1996–98 both p and V also refer to those employees who work unpaid overtime and/or who are compensated by leisure. Notice that p and V are only observed from 1996 onwards. By construction, $p = 0$ is necessary and sufficient for $V = 0$. We identify three types of plant:

$D = 0$: plants where every worker works zero overtime ($p = 0$)
$D = 1$: plants where a proportion of workers work positive overtime ($0 < p < 1$)
$D = 2$: plants where every worker works positive overtime ($p = 1$)

Slightly less than half of the IAB dataset (45%) comprises overtime plants ($D = 1, 2$), of which very few plants (10%) employ entirely overtime workers ($D = 2$). The average proportion of overtime workers in overtime plants is 41%. When weighted, these numbers—given in the third column from the right of Table 3—change a lot: 25%, 28%, 58%. The corresponding unweighted numbers for the regression samples are given in the two rightmost columns of Table 3.

Of the remaining variables, there is no information on either $\widehat{z}$ or $\bar{z}$. For the hourly wage rates $\widehat{w}$ and $\bar{w}$, all we observe is total labour costs in the plant:

$$C = w(\bar{H} + \gamma V)\widehat{N} + w\bar{H}\bar{N}, \quad \text{assuming } w = \bar{w} = \widehat{w}, \tag{13}$$

from which an hourly wage rate w can be computed using $\gamma = 1.25$.[18]

Three other key variables are listed in Table 3, which also reports weighted sample means, by the two plant-types, and by year. Two are included to control for the scale of the plant. Received wisdom is that the capital stock K should be used, which is a much better control than real output Y which is jointly determined with employment and hours. Unfortunately, we only observe investment I, which is of some use in regressions

[17]This was modified, in 1998, to "How long is the average currently agreed weekly working time for full-time workers?", to take account of flexible bargaining agreements, which allow $\bar{H}$ to vary over the workforce.

[18]In Germany, most bargaining agreements fix γ at 1.25 (as obtained from the WSI-Tarifarchiv, BMA-Tarifregister); however, γ is typically larger for weekend or night work.

where the change in K is used. We also model aggregate demand θ using a production index for the plant's two-digit industry. The final variable is a union bargaining dummy B for whether there is an agreement at either the plant– or industry–level.

Some other facts in Table 3 are noteworthy. Overtime plants are bigger, on average, than standard-time plants. They also have a higher standard workweek, in spite of offering overtime. Also, average overtime is about three hours, worked by about 60% of employees on average. There are also some differences between East and West. Overtime plants are bigger in the West whereas standard-time plants are about the same size. The standard workweek is one hour shorter in the West and the hourly wage rate 2.5 *DM* higher, giving workers in the West a higher weekly income. Also, labour productivity is nearly twice as high in the West, while investment per head is 1.6 times larger in the East. For these reasons, we stratify by East/West throughout our empirical work.

Average employment has fallen in both East and West. At the same time, the proportion of overtime plants has fallen considerably, from 30% to 22% in the West (over 6 years) and from 29% to 19% in the East (over 4 years). (See the row for D in Table 3.) In both the West and East, standard hours fell very little between 1995 and 1999. This is the first time in many years that, on average, standard hours have not fallen.[19] Estimates of worksharing effects in the employment regressions are driven by cross-section correlations between changes in standard hours and changes in employment. In the IAB dataset, 80% of all plant-year observations saw no change in standard hours; of those that did, 63% were cuts whereas 37% were increases (Schank 2001, Table 4.6). A bigger proportion of plants changed $\bar{H}$ in the early part of the sample: in the West the proportion was 27% for 1993/95 but dropping to 20% in later years; for the East, the corresponding numbers for 1996 are 32% and 15%. Finally, the variance of standard hours, about a mode of 40 hours, is much smaller in the East (see Figure 3).

Table 4 reports the basic regression samples. For regressions with employment as the dependent variable (1993–99) there are 18,596 plant-years; for regressions with the proportion of overtime workers as the dependent variable ('employment decomposition regressions') there are 13,163 plant-years. The latter sample size is smaller because p is only observed for 1996–99, and, in addition, there are missing values for p. The corresponding number of plants are also reported in the table. The number of differences, that is the number of plant-years corresponding to plants that have two or more observations, is obviously smaller.

[19] See Bauer & Zimmermann (1999, Figure 4) and Hunt (1999).

5 Employment Regressions

5.1 Econometric issues and basic results

Below we report estimates of standard labour demand functions which include $\bar{H}$ as the covariate of interest, recognising that its effect (and other variables) may depend on whether the observation refers to an overtime or standard-time plant. Because the wage is often endogenous for many plants, being negotiated over with unions, we estimate reduced-form labour demand equations where the wage is absent. The estimated employment-standard hours elasticity is of more policy interest than those in Equations (4, 5) above.

One specific issue we consider throughtout is the extent to which there is heterogeneity in any estimated worksharing effects. We therefore stratify pairwise between East and West and between the agriculture/manufacturing and service sectors. (Hereafter we refer to the former as 'manufacturing'.) As noted already, during the sample period, there were big differences between East and West German economies; the East had lower productivity, lower wages, fewer unionised plants, a longer and less variable standard working week, and more investment. Also, unemployment was higher, and so it might be easier to substitute unemployed workers for hours. We stratify between the two sectors because they have different production processes and because unpaid overtime is more prevalent in the service sector.

Unobserved scale effects

The perennial problem with estimating labour demand functions is that very spurious effects will be estimated unless the scale of the plant is controlled for. Suppose that the true labour demand model is:

$$n_{it} = \eta \bar{h}_{it} + \mathbf{x}'_{it}\boldsymbol{\beta} + \gamma k_{it} + f_i + u_{it}, \tag{14}$$

where $n \equiv \log N$, $\bar{h} \equiv \log \bar{H}$, $\mathbf{x}'_{it}$ is a vector of other covariates, and f_i is a plant-specific time-invariant unobservable, which may be modelled as either a random or fixed effect, depending on whether it is assumed uncorrelated with the observed covariates. k_{it} is the log of the scale of the plant and, because of constant returns to scale, we expect $\gamma \approx 1$.

If k_{it} is not observed, then OLS or random effects (GLS) estimates of η will be downwards biased (too negative), because the bias depends on the correlation between $\bar{h}_{it}$ and k_{it} and the true value of γ. It seems reasonable to assume that the correlation between standard hours and scale is negative as large plants are typically more unionised, with bargains resulting in lower standard hours than in the non-union sector. (See Andrews & Simmons (2001) for further evidence and a bargaining model with such a prediction.) Moreover, the size of the bias is going to be considerable, given $\gamma \approx 1$.

If k_{it} is inadequately proxied by observables, then part of the scale effect will be picked up by f_i, particularly if the underlying scale effect is time-invariant. Of course, f_i itself may well be correlated with $\bar{H}$. The standard way to test for the correlation between f_i and all the covariates is a Hausman test, computed by comparing random effects (GLS) and fixed effects (covariance) estimates. Throughout, the null of no correlation is massively rejected and, as expected, implausible estimates of η (large negative) are obtained, and are, in fact, quite close to their OLS counterparts.

The decision whether to estimate the model in first differences or use the covariance transform depends on which give the more efficient estimates. Both estimators are consistent. Following Wooldridge (2002, Section 10.6.3), we estimate the model in first differences,

$$\Delta n_{it} = \eta \Delta \bar{h}_{it} + \Delta \mathbf{x}'_{it} \boldsymbol{\beta} + \gamma \Delta k_{it} + \Delta u_{it}, \tag{15}$$

and then test whether the residuals $\widehat{\Delta} u_{it}$ exhibit first-order serial correlation by regressing them on $\widehat{\Delta} u_{i,t-1}$. A serial correlation estimate approaching -0.5 suggests that the within estimator is more efficient.

The first set of estimates are reported in Table 5 as 'BaseFD', together with Wooldridge's test, denoted 'SC test'. Also note that Δk_{it} is proxied by investment (using output—in fact, total sales—differences, appropriately instrumented, made little difference). Also included in $\mathbf{x}'_{it}$ are $T-1$ differenced time-dummies, where $T = 6$ years.

The worksharing estimates themselves are discussed more fully below; however, there is a sizeable, significant negative estimate of $\eta = -0.754$ for manufacturing in East Germany. The effect in the East Service sector is also negative, but smaller and insignificant. The effects for both West German sectors are zero, as it is for the whole sample.

Wooldridge's serial correlation test is either insignificant or the estimate is much closer to zero than -0.5. We only report differenced estimates, therefore. The price of estimating the model in differences (or mean-deviations) are potentially threefold. First, the effects of other covariates are either swept away or rely solely on time-series (within) variation in the data. However, these estimates are not of interest. Second, measurement error means estimates are too close to zero, but we argue that the wording of the question on standard hours rules out this possibility *a priori*. Third, there needs to be sufficient variation in $\Delta \bar{h}_{it}$ across plants. As already noted above, enough plants have changed standard hours in the sample; our estimates are effectively based on 2,851 non-zero differences (again see Table 4).

Weighting

Given the highly variable weights discussed above, here we discuss whether or not we should use weighted or unweighted estimation, correctly incorporated into a model with fixed effects. The issue is whether it is a single worksharing parameter in the population that is being modelled, or whether there is likely to be heterogeneity in the worksharing effect across various strata s. When there is heterogeneity, the classic argument is that both OLS and weighted least squares (WLS) are inconsistent, but if there is no heterogeneity, OLS will be more efficient (because of Gauss-Markov) (Deaton 1997). Deaton recommends the computation of a Hausman-type test that compares the two sets of parameter estimates. If the null hypothesis is not rejected, one would use OLS estimates; otherwise, further modelling of the heterogeneity is needed. This is why we stratify the sample between East and West Germany and between the manufacturing and service sectors, and why we also examine below whether worksharing varies between standard time and overtime plants, between union and non-union plants and across plant size.

The appropriate weighted least squares (WLS) regression that sweeps out the fixed effects using a covariance transform is to use deviations from the probability weighted means across time for each plant, and then use GLS, using ω as a weight. This makes intuitive sense in that, in the absence of weights, one would use deviations from unweighted time-series means. For any variable z, this means computing $z_{it} - \sum_t \omega_{it} z_{it} / \sum_t \omega_{it}$. When $T_i = 2$, this regression is the same as using differences, but using $\omega_{i1}\omega_{i2}/(\omega_{i1}+\omega_{i2})$ as weights. Thus Equation (15) is estimated using $\omega_{it}\omega_{i,t-1}/(\omega_{it}+\omega_{i,t-1})$ as weights. Notice that there is no requirement that the weights have to vary across time. Clearly this regression can be compared with Equation (15) without weights (ie set to unity). Given we are only interested in the effects of standard hours, we only compute a Hausman test for the equality between the parameter(s) on standard hours.

The WLS results are reported in the rows denoted 'FD weighted', with the Hausman test immediately below. The estimates are very similar to 'FD unweighted' and consequently the Hausman test is not rejected easily in all five cases (the estimate for East manufacturing moves from -0.754 to -0.914). Given this, and that modern practice says that its is better to use OLS estimates, but correcting the covariance matrix for arbitrary heteroskedasticity, hereafter we dispense with WLS estimates. The covariance matrix is clustered using the 320 strata s.

Are standard hours endogenous?

We now consider whether is it legitimate to assume that standard hours are strictly exogenous, $\mathrm{E}(u_{it}|\bar{h}_{it}, f_i) = 0$. This assumption partly depends on the 'level' at which bargaining with unions takes place. In Germany, if the plant is unionised, bargaining over standard hours usually takes place at the industry, regional, or even national level. Recall that, historically, most changes in standard hours have been driven by two large unions (IG Metall and IG Druck). Here, one can safely assume that the plant is on its labour demand schedule. Similar considerations apply to plants who do not bargain with unions and who operate in a competitive labour market. In both cases, one might argue that standard hours are strictly exogenous. On the other hand, if union-plant negotiations are at the plant-level, one might argue that standard hours are endogenous (and the issue of worksharing might be part of such negotiations). For example, in Andrews & Simmons's (2001) model, standard hours and the hourly wage rate are on a downward-sloping contract curve located to the right of the hours demand schedule, which suggests that $\bar{h}_{it}$ needs an instrument. Before we discuss such an instrument, there are other reasons why $\bar{h}_{it}$ and u_{it} might be correlated, and so we proceed first by testing for strict exogeneity following Wooldridge (2002, Section 10.7.1). Here we add the level $\bar{h}_{it}$ to Equation (15) and test for its significance: see 'Endogeneity test' in Table 5. Although it is sometimes significant, in practical terms the worksharing estimates are unaffected.

Our candidate instrument is industry-level standard hours, $\bar{h}^*_{it} \equiv \log \bar{H}^*_{it}$, which should be uncorrelated with u_{it}, but will in general have some influence on negotiated hours; that is, we can check the extent to which $\bar{h}^*_{it}$ and $\bar{h}_{it}$ are correlated. We therefore report IV estimates of Equation (15) using the following instrument,

$$1\{B = 0, 1\}\bar{h} + 1\{B = 2\}\bar{h}^*,$$

as a difference, where $B = 0$ indicates that a plant does not have a bargaining agreement, $B = 1$ indicates that a plant has a bargaining agreement at the industry level, $B = 2$ indicates that a plant has a bargaining agreement at the plant-level, and the function $1\{\ \}$ takes the value unity if the statement in $\{\ \}$ is true, zero otherwise. This assumes that only plant-level bargains are endogenous. Because plants may have some market power even when they do not negotiate with unions, we also use, as an alternative:

$$1\{B = 1\}\bar{h} + 1\{B = 0, 2\}\bar{h}^*.$$

The instrumental estimates are given under 'FDIV' for both variants. When we use the first instrument, very little changes, mainly because plant-level bargains represent only 4.7% of the sample. For the second instrument, the estimates do change, with the estimate for East/manufacturing moving from -0.754 to -0.839. However, the standard error doubles, because the partial correlation between $\bar{h}$ and $\bar{h}^*$ is 0.679 (in levels, 0.830) (Wooldridge 1999, Eqn. 15.13). See the row labelled 'First Stage FDIV'. However, notice that the sample sizes are smaller because data on $\bar{h}^*$ are not available for all plants, and it is possible that the estimated effect on all the data might have been stronger (in fact, the BaseFD estimate falls from -0.754 to -0.511 with 374 fewer observations).

The main result from Table 5—and the most important result of the paper thus far—is that there is a strong negative (pro-worksharing) effect in the East manufacturing sector (elasticity of -0.839). In other words, a 2-hour reduction in normal working time (ie a 5% reduction) increases employment on average by about 4% in this sector. There are no worksharing effects estimated for West Germany. Because the FDIV and FD (weighted and unweighted) estimates are very similar, we stick with unweighted OLS in what follows. In particular, instrumenting standard hours in censored regression models later on is problematic.

5.2 Other specifications

We first investigate whether worksharing varies between standard-time and overtime plants. The only other investigation of this issue (Hart & Wilson 1988) recognised that the theory predicts that worksharing is less likely in overtime plants (see Subsections 3.1 and 3.2 and especially Figure 2) and so we group together all plants that have at least some overtime working into one category, ie $D_{it} = 1, 2$. In Equation (14), we replace $\eta \bar{h}_{it}$ by

$$\eta_0 1\{D_{it} = 0\}\bar{h}_{it} + \eta_1 1\{D_{it} = 1, 2\}\bar{h}_{it} + \eta_2 1\{D_{it} = 1, 2\},$$

and then estimate in differences. We do not report these regressions as the t-test for the hypothesis that $\eta_0 = \eta_1$ is rejected in all 5 cases. In other words, the estimates of η do not depend on working-time regime ($D = 0$ versus $D = 1, 2$). We also investigated the possibility that plants who change status might have a different effect to those who do not change, but found nothing significant. This suggests that the discontinuous behaviour implied by Figure 2 is not found in the data, and that firms are more likely to adjust to changes in standard hours by offering/withdrawing overtime to only a proportion of its workforce.

We next investigate whether worksharing effects vary by the bargaining dummy B, three plant-size categories, and whether the plant was born before unification in 1990. The next table (6) reports variations by plant-size, for both FD and FDIV estimators. In other words, we replace $\eta \bar{h}_{it}$ by

$$\eta_3 1\{N_{it} < 15\}\bar{h}_{it} + \eta_4 1\{15 \leq N_{it} < 100\}\bar{h}_{it} + \eta_5 1\{N_{it} \geq 100\}\bar{h}_{it} + \eta_6 1\{B_{it} = 1, 2\}\bar{h}_{it},$$

and then estimate in differences. The table suggests that the strong worksharing effects detected in East German manufacturing are driven by small plants (those employing less than 15 employees), although there are also negative effects estimated for the other size categories, although insignificant. There is no evidence that the presence of a bargaining agreement affects worksharing (reported), nor whether the plant is older than 1990 (not reported). We now attempt to explain *why* worksharing occurs only in small East German plants.

5.3 Why is there worksharing in small, East German manufacturing plants?

It is the case that East German plants are different from West German plants, as already noted. Establishing why this means differences in worksharing effects is a lot harder. In what follows, we list a number of reasons, and provide evidence, if available in our data.

All of the regressions reported thus far have employment as the dependent variable. We also re-estimate all of the models reported in Tables 5 and 6, but with the hourly wage rate and total weekly hours on the left-hand-side. These are not reported in the tables. When modelling total weekly hours, the Stylised Fact on Page 7 is just as true for our plant-level data. Virtually every estimate obtained (weighted/unweighted, OLS/IV, fixed/random effects, with/without censoring) lies within two standard errors of unity. In other words, changing standard hours has no effect on overtime. However, there are no differences for small, East German, manufacturing plants.

The regressions for hourly wage rates are more interesting. It is generally accepted that the way to interpret a negative effect of standard hours on wages is as a supply side response: workers negotiate for higher hourly wages (so-called wage compensation). If the labour demand model includes the hourly wage rate as a covariate (recall, ours do not), this adds a negative indirect impact on employment in addition to any direct effect from standard hours. These effects can be sizeable if the labour demand elasticity for wages is large, or the effect of standard hours on wage rates is large. The latter is often the case: an elasticity of minus unity means that workers' incomes are unaffacted by cuts in standard hours.

Our wage regressions reveal that income compensation is lowest for small manufacturing plants in East Germany, with a *positive* elasticity of 0.117. The other estimates are -0.321 for West/manufacturing, -1.36 for West/service and -0.609 or East/service, giving an overall estimate of -0.905 for all small plants. In general, wage costs are lower in East Germany, where employers might be more willing to increase employment after a cut in standard hours, although this clearly did not happen in June 2003 (see the Introduction). This provides one possible, fairly convincing, reason for the pro-worksharing effects being discussed.

One of the standard predictions of the theory is that worksharing is less likely where quasi-fixed labour costs are higher. Whilst this might explain why small plants might adjust employment rather than hours, it does not explain differences between East and West. However, in East Germany, it is likely that small plants might be more efficient and flexible than larger ones, as they tend to be younger and less likely to suffer from working practices inherited from before re-unification. We attempted to find evidence for this by adding a dummy for whether a plant was born before 1990, but this revealed nothing as it is so highly correlated with firm size.

Next, there are lots of employment programmes in East Germany where plants get subsidies for hiring employees. As these plants increase their workforce, it is possible they lower their standard working time. Although the IAB-panel includes information on employment subsidies, this information is basically a fixed-effect and is therefore useless in panel estimations.

Finally, the concept of a standard working is more relevant for blue-collar compared with white-collar workers; the latter, if working overtime, will work unpaid overtime, for which the theory does not apply. Since the proportion of blue-collar workers is much larger in manufacturing than within services, this may explain why we detect worksharing effects in the East, especially manufacturing.

To conclude, we suspect that there is no single explanation for why we find worksharing effects in small East German manufacturing plants. It is also worth noting that these effects might disappear through time as East Germany converges towards the West.

6 Employment decomposition regressions

The next set of regressions decompose the impact of standard hours ($\bar{H}$) on employment (N) into the impact on the number of overtime workers ($\widehat{N}$) and the number of standard-time workers ($\bar{N}$), thereby testing the model of Section 3.3, which predicts that the proportion of overtime workers (p) falls after a cut in the normal working time. We do this by estimating models for N (as above) and p:

$$N_{it} = a_1\bar{h}_{it} + \mathbf{x}'_{it}\boldsymbol{\beta}_1 + \gamma_1 k_{it} + f_i + v_{it} \tag{16}$$

$$p^*_{it} = a_2\bar{h}_{it} + \mathbf{x}'_{it}\boldsymbol{\beta}_2 + \gamma_2 k_{it} + \xi_i + \varepsilon_{it} \tag{17}$$

$$p_{it} = \begin{cases} 0 & \text{if } p^*_{it} \leq 0 \ (D_{it} = 0), \text{ i.e. } N_{it} = \bar{N}_{it} \\ 1 & \text{if } p^*_{it} \geq 0 \ (D_{it} = 2), \text{ i.e. } N_{it} = \widehat{N}_{it} \\ p^*_{it} & \text{else } (D_{it} = 1), \text{ i.e. } N_{it} = \bar{N}_{it} + \widehat{N}_{it} \end{cases}$$

$$\bar{N}_{it} \equiv (1 - p_{it})N_{it}; \qquad \widehat{N}_{it} \equiv p_{it}N_{it} \tag{18}$$

The parameter a_1 is different to η in Equation (14) above, as the dependent variable is no longer in logarithms. As above, Equation (16) is estimated by differencing out the fixed-effects.

Equation (17) is more succinctly written as $p^* = \mathbf{z}'\boldsymbol{\delta} + \varepsilon$. Assuming that ε is Normally distributed, Equation (17) is estimated as a two-sided censored regression (hereafter referred to as a two-sided Tobit). Figure 4 suggests that Normality is a reasonable assumption, at least for $0 < p_{it} < 1$. ξ_i is a fixed-effect, just like f_i, and is again potentially correlated with $\bar{h}_{it}$. Estimation of Tobit-type fixed-effects models is less straightforward than for Poisson- or Logit-type models (see Arellano & Honoré (2001) for a recent survey). However, Wooldridge (2002, Section 16.8.2) recommends the following under the assumption of strict exogeneity, one we have made throughout. The potential correlation between standard hours and ξ_i is modelled by $\xi_i = a_3\bar{\bar{h}}_i + \omega_i$, where $\bar{\bar{h}}_i$ is average standard hours for plant i and ω_i is a random effect. Substituting into Equation (17) suggests that the two-sided Tobit be estimated as a random effects model—straightforward in Stata, for example—but with $\bar{\bar{h}}_i$ as an extra covariate.

It turns out that $\bar{\bar{h}}_i$ is insignificant in all 5 regressions; the p-values being 0.37, 0.34, 0.58, 0.34 and 0.14. The reason why ξ_i is uncorrelated with average standard hours is probably because there are no scale effects in the model for p^*. This is because p^* is a logit function of $\log \bar{N} - \log \widehat{N}$, and so scale effects approximately 'difference out'. In what follows we dispense with the efficiency gains by estimating the model, without $\bar{\bar{h}}_i$, as a random effects two-sided Tobit because this would complicate what follows later.

The parameter of interest is a_2, but the marginal effect of p with respect to standard hours is also computed (with $\mathbf{z}'\boldsymbol{\delta}$ evaluated at the sample means of the data):

$$\frac{\partial p}{\partial \bar{h}} = \left[\Phi\left(\frac{1 - \mathbf{z}'\boldsymbol{\delta}}{\sigma}\right) - \Phi\left(\frac{-\mathbf{z}'\boldsymbol{\delta}}{\sigma}\right)\right] a_2. \tag{19}$$

The term in square brackets is a positive fraction.[20] The effect of standard hours on

[20] More correctly written as $\frac{\partial \mathrm{E}(p)}{\partial \bar{h}}$. Similarly, the marginal effect for the censored observations, $\frac{\partial \mathrm{E}(p|0<p<1)}{\partial \bar{h}}$, can also be computed, but this is just another positive fraction of a_2.

plant-type is also computed:

$$\begin{aligned}
\frac{\partial \Pr(D=0)}{\partial \bar{h}} &= -\phi\left(\frac{-\mathbf{z}'\boldsymbol{\delta}}{\sigma}\right)\frac{a_2}{\sigma} < 0 \\
\frac{\partial \Pr(D=2)}{\partial \bar{h}} &= \phi\left(\frac{1-\mathbf{z}'\boldsymbol{\delta}}{\sigma}\right)\frac{a_2}{\sigma} > 0 \\
\frac{\partial \Pr(D=1)}{\partial \bar{h}} &= \left[\phi\left(\frac{-\mathbf{z}'\boldsymbol{\delta}}{\sigma}\right) - \phi\left(\frac{1-\mathbf{z}'\boldsymbol{\delta}}{\sigma}\right)\right]\frac{a_2}{\sigma} \gtrless 0.
\end{aligned} \tag{20}$$

It is easy to see that a cut in standard hours shifts the distribution of p^* to the left (assuming $a_2 > 0$), generating a larger proportion of the sample as left censored, but a smaller proportion as right censored. Thus the proportion of overtime plants ($D = 1, 2$) should fall, which is at odds with the small amount of evidence we have from individual-level data (Table 1). This result is guaranteed if $a_2 > 0$. It is also a different prediction to the models where plants employ either entirely standard-time workers or entirely overtime workers, as illustrated in Figure 2.

Equations (16) and (17) are estimated on the same sample (1996–99 only, since p is only observed for these years) and our results are presented in Table 7. The results for the employment equations are very similar to those estimated on the full sample (1993–99), that is, the finding that pro-worksharing effects are confined to the manufacturing sector of East Germany still holds up.

Turning to the five equivalent equations for the proportion of overtime workers (p), estimated as two-sided Tobit regressions, first notice that the parameter associated with standard hours, $a_2 \equiv \partial p^*/\partial \bar{h}$, is estimated at 0.35. This is approximately equal to the weighted average of the four sector estimates, being zero for both manufacturing sectors and about 0.7 for both service sectors. These two positive effects are exactly as predicted by the theory. For the West German service sector, a cut in standard hours of 4 hours (10%) leads to a 4.59 percentage point increase in the number of standard-time plants ($D = 0$), with the $D = 1$ plants falling by about the same amount. ($D = 2$ plants are a very small proportion of total.) These are large semi-elasticities. Notice that the 0.71 estimate converts to an estimate on the *observed* proportion of overtime workers employed in the plant as $\partial p/\partial \bar{h} = 0.23$. The reason for this smaller impact is that, as with all censored regression models, as standard hours fall, some plants substitute completely away from employing overtime workers and therefore change their working-time regime to a standard-time plant.

Although the aim of the exercise is estimate $\partial p/\partial \bar{h}$, it is interesting to compute $\eta_{\hat{N}\bar{H}}$ and and $\eta_{\bar{N}\bar{H}}$, given estimates of the elasticity $\eta_{N\bar{H}}$ from Equation (16) and the semi-elasticity $\partial p/\partial \bar{h}$ from Equation (17). From Equation (18):

$$\frac{1}{\widehat{N}_{it}}\frac{\partial \widehat{N}}{\partial \bar{h}} = \frac{1}{p_{it}}\frac{\partial p}{\partial \bar{h}} + \frac{1}{N_{it}}\frac{\partial N}{\partial \bar{h}} \quad \text{or} \quad \eta_{\widehat{N}\bar{H}} \equiv \frac{1}{p_{it}}\frac{\partial p}{\partial \bar{h}} + \eta_{N\bar{H}} \quad \text{if } 0 < p_{it} \leq 1 \tag{21}$$

$$\frac{1}{\bar{N}_{it}}\frac{\partial \bar{N}}{\partial \bar{h}} = -\frac{1}{1-p_{it}}\frac{\partial p}{\partial \bar{h}} + \frac{1}{N_{it}}\frac{\partial N}{\partial \bar{h}} \quad \text{or} \quad \eta_{\bar{N}\bar{H}} \equiv -\frac{1}{1-p_{it}}\frac{\partial p}{\partial \bar{h}} + \eta_{N\bar{H}} \quad \text{if } 0 \leq p_{it} < 1. \tag{22}$$

Thus we report in Table 7 that $\eta_{\bar{N}\bar{H}} = -0.238$ (ranging from zero to -0.845) and $\eta_{\widehat{N}\bar{H}} = 0.199$ (ranging from -0.605 to 0.230), where we have evaluated p_{it} at the appropriate sample average of p (also given in Table 7). Of course, it is no surprise that these estimates are sensible, given the estimate of $\partial N/\partial \bar{h}$ and given that $\partial p^*/\partial \bar{h}$ and therefore $\partial p/\partial \bar{h}$ are estimated as positive. In effect, we have decomposed the total worksharing elasticity $\eta_{N\bar{H}}$ into elasticities for the two types of worker, in so far as $\eta_{N\bar{H}}$ is approximately equal to $\eta_{\widehat{N}\bar{H}} + \eta_{\bar{N}\bar{H}}$ for three of the five columns.

The objective of this section was to use data on p to see whether the theory is correct in predicting that a cut in standard hours leads to more standard time workers, but fewer overtime workers. The final panel of Table 7 clearly illustrates this substitution effect for the two service sectors. However, the theory is silent on whether or not total employment increases (worksharing)—the evidence suggests that it remains constant except for small plants in East German manufacturing.

7 Conclusion

In this paper, we report new estimates of the effect of changing standard hours on employment, using a panel of German plants (the IAB panel) for the period 1993–99. Hitherto there has been virtually no microeconometric evidence with which to assess this issue, with Crépon & Kramarz (2002) being a notable exception. In our data, not only are we able to observe whether a plant is an overtime plant or standard-time plant, but also we observe the proportion of overtime workers employed in the plant.

We develop a compositional model to incorporate these two variables. Our model shows that worksharing is basically a substitution effect, whereby a cut in standard hours leads to a smaller proportion of overtime workers. It is silent as to whether total employment increases; in other words, it is an empirical issue as to whether worksharing works.

Although the data over-sample large firms, our analysis has checked that unweighted methods are appropriate. We also use differences to sweep out unobserved fixed effects (primarily caused by not observing the scale of the plant), and use IV methods to control for the potential endogeneity of standard hours. We stratify across four sectors, that is East and West Germany taken pairwise with manufacturing/agriculture and service sectors.

Our results are as follows:

1. Apart from one exception, there is no evidence of worksharing with these data. The presence of unions has no effect, nor does the working-time regime of the plant. This contrasts with Crépon & Kramarz (2002), but our study is not really comparable, using plant-level rather than individual-level data. Also, Crépon & Kramarz (2002) analyse data which are much more like a natural experiment than in Germany, where cuts in standard hours occur frequently, but are not part of a single one-off policy change.

2. There are large pro-worksharing effects in small plants (fewer than 100 employees) in the East, non-service sector ($\eta_{N\bar{H}} \approx -3/4$). We discuss various reasons why this is so, and conclude that these arise because the East German non-service sector is different from the West. As these differences will disappear with convergence, we predict that this worksharing result will weaken through time.

3. Because we observe the proportion of overtime workers employed in the plant, we use a two-sided Tobit regression to estimate the effect of standard hours. Our estimate $\partial p^*/\partial\bar{h}$ is about 0.7 in the two non-service sectors, but zero otherwise. We then (approximately) decompose the effect of standard-time changes on employment $\eta_{N\bar{H}}$ into separate effects for overtime workers $\eta_{\widehat{N}\bar{H}}$ and standard-time workers $\eta_{\bar{N}\bar{H}}$. These estimates vary across the four sectors, depending on the estimates of $\partial p^*/\partial\bar{h}$ and $\partial \log N/\partial\bar{h}$.

4. Because we also observe the plant's working-time regime, the two-sided Tobit suggests that $\partial Pr(D=0)/\partial\bar{h}$ is estimated as approximately –1/4. This estimate for the impact of standard hours on the proportion of standard-time plants is consistent with the theory, but is completely at odds with all the evidence from individual-level surveys. This might explain why our result is different from Crépon & Kramarz's, although one would need matched employee-employer data to resolve this issue.

To conclude, in our compositional model of demand for hours and employment, we explicitly model the proportion of workers performing overtime in plants. This generates a richer model of demand for hours and employment than has been available hitherto and permits a rigourous empirical treatment of the worksharing issue when applied to the German IAB establishment-level panel.

Appendix A Comparative statics

A.1 Overtime firms

See Section 3.1. An interesting special case of 'no scale effects' is when $R(g(H)N) = (g(H)N)^\beta$ and $g(H) = H^{\alpha/\beta}$, $0 < \alpha < \beta < 1$, which generates the familiar Cobb-Douglas $\theta H^\alpha N^\beta$. Tractable and realistic demand functions can then be derived:

$$H = \frac{\alpha}{\gamma(\beta-\alpha)}\frac{z}{w} - \frac{\alpha(\gamma-1)}{\gamma(\beta-\alpha)}\bar{H} \tag{A.1}$$

$$N = \left(\frac{\gamma}{\alpha}\frac{w}{\theta}\right)^{\frac{1}{\beta-1}} H^{\frac{1-\alpha}{\beta-1}} \qquad \text{with } H \text{ as above.} \tag{A.2}$$

It is self-evident that $\bar{H}$ cannot be too high relative to the values of other exogenous variables in the model or else the constraint given by Equation (6) is violated because the firm demands too few hours. It is easy to show that the necessary condition is $\bar{H} \leq [\alpha/(\gamma\beta-\alpha)](z/w)$, which can be rewritten as:

$$k \equiv \frac{z}{w\bar{H}+z} \geq 1 - \frac{\alpha}{\gamma\beta}. \tag{A.3}$$

In other words, the share of fixed costs in 'normal' costs must exceed the right-hand-side of the equation. Plausible values for these parameters ($\beta = 3/5$, $\alpha = 2/5$, $\gamma = 3/2$) would imply that $k \geq 5/9$ is much higher than the one-quarter figure often quoted (Hart 1984).

A.2 Standard-time firms

The Cobb-Douglas case is illustrative:

$$N = \left(\frac{w\bar{H}+z}{\theta\beta\bar{H}^\alpha}\right)^{\frac{1}{\beta-1}}. \tag{A.4}$$

It is easy to show that $\eta_{N\bar{H}}$ is given by:

$$\eta_{N\bar{H}} = \frac{1-k-\alpha}{\beta-1} \leq 1 - \frac{1}{\gamma\beta}$$

with the inequality coming from (A.3). Worksharing works if $\alpha < 1-k$. It is possible to find parameter value so that $\eta_{N\bar{H}} > 0$ (e.g. $\beta = 4/5$, $\alpha = 1/5$, $\gamma = 3/2$, $k = 0.82$) but the set is small, and the values of k needed are implausibly large.

A.3 Two types of worker within a firm

See Section 3.3. Comparative statics come from differentiating Equations (8–10). (In this subsection, we set $\theta = 1$.)

$$\begin{bmatrix} R_{\widehat{N}\widehat{N}} & R_{\widehat{N}\bar{N}} & R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} \\ & R_{\bar{N}\bar{N}} & R_{\bar{N}\widehat{H}} \\ & & R_{\widehat{H}\widehat{H}} \end{bmatrix} \begin{bmatrix} d\widehat{N} \\ d\bar{N} \\ d\widehat{H} \end{bmatrix} = \begin{bmatrix} 0 \\ \bar{w} \\ 0 \end{bmatrix} d\bar{H} + \begin{bmatrix} -(\gamma-1)\widehat{w} \\ 0 \\ 0 \end{bmatrix} d\bar{H} - \begin{bmatrix} R_{\widehat{N}\bar{H}} \\ R_{\bar{N}\bar{H}} \\ R_{\widehat{H}\bar{H}} \end{bmatrix} d\bar{H}. \quad \text{(A.5)}$$

Concavity implies that $R_{\widehat{N}\widehat{N}} < 0$, $R_{\bar{N}\bar{N}} < 0$, $R_{\widehat{H}\widehat{H}} < 0$, and $|\mathbf{R}| < 0$. As usual, we also assume that $R_{\widehat{N}\bar{N}} > 0$, $R_{\bar{N}\widehat{N}} > 0$, and $R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} < 0$. Each term on the RHS of the equation corresponds to the 3 'channels' discussed in the main text. To establish the effect of standard hours $\bar{H}$ on the proportion of overtime workers p, note that $p = \Lambda(\bar{N}/\widehat{N})$, where $\Lambda()$ is the Logit function, and so

$$\frac{\partial p}{\partial \bar{H}} = \Lambda' \frac{\bar{N}}{\widehat{N}} \left(\frac{1}{\bar{N}} \frac{\partial \bar{N}}{\partial \bar{H}} - \frac{1}{\widehat{N}} \frac{\partial \widehat{N}}{\partial \bar{H}} \right), \qquad \Lambda' > 0.$$

For the first channel, it follows that, for any revenue function:

$$\frac{\partial \bar{N}}{\partial \bar{H}} = \frac{\bar{w}}{|\mathbf{R}|} \begin{vmatrix} R_{\widehat{N}\widehat{N}} & R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} \\ R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} & R_{\widehat{H}\widehat{H}} \end{vmatrix} < 0$$

$$\frac{\partial \widehat{N}}{\partial \bar{H}} = \frac{\bar{w}}{|\mathbf{R}|} \begin{vmatrix} R_{\widehat{N}\bar{N}} & R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} \\ R_{\widehat{H}\bar{N}} & R_{\widehat{H}\widehat{H}} \end{vmatrix} \lesseqgtr 0$$

which is the usual negative own-price effect and ambiguous cross-price effect, the latter depending on the balance of substitution and scale effects. For the second channel, the price of $\widehat{N}$ is $-\widehat{w}(\gamma-1)$, and so

$$\frac{\partial \widehat{N}}{\partial \bar{H}} = \frac{-\widehat{w}(\gamma-1)}{|\mathbf{R}|} \begin{vmatrix} R_{\bar{N}\bar{N}} & R_{\bar{N}\widehat{H}} \\ R_{\widehat{H}\bar{N}} & R_{\widehat{H}\widehat{H}} \end{vmatrix} > 0$$

$$\frac{\partial \bar{N}}{\partial \bar{H}} = \frac{\widehat{w}(\gamma-1)}{|\mathbf{R}|} \begin{vmatrix} R_{\bar{N}\widehat{N}} & R_{\bar{N}\widehat{H}} \\ R_{\widehat{N}\widehat{H}} - R_{\widehat{H}}/\widehat{N} & R_{\widehat{H}\widehat{H}} \end{vmatrix} \lesseqgtr 0.$$

Clearly, the effect on employment is ambiguous because the scale effects operate in different directions.

To show that the substitution effects $\partial p/\partial \bar{H}$ are unambiguously positive, it is necessary to use the no-scale-effect revenue function, whereby: $R_{\widehat{N}\widehat{N}} = R_{LL}f^2 < 0$, $R_{\widehat{N}\bar{N}} = R_{GL}fg > 0$, $R_{\widehat{N}\widehat{H}} = R_{LL}ff'\widehat{N} + R_L f' < 0$, $R_{\bar{N}\bar{N}} = R_{GG}g^2 < 0$, $R_{\bar{N}\widehat{H}} = R_{GL}gf'\widehat{N} > 0$, $R_{\widehat{H}\widehat{H}} = R_{LL}f'f'\widehat{N}^2 + R_L f''\widehat{N} < 0$, and

$$|\mathbf{R}| = (R_{GG}R_{LL} - R_{GL}^2)f^2g^2f''\widehat{N} < 0.$$

It follows that, adding both channels,

$$\begin{aligned}\frac{\partial p}{\partial \bar{H}} = -\Lambda'\widehat{w}(R_{GG}R_{LL} - R_{GL}^2)g^2f'f'\widehat{N}/\left|\mathbf{R}\right| - \Lambda'\widehat{w}(R_{GG}g/\widehat{N} + R_{GL}f/\bar{N})gR_Lf''\widehat{N}/\left|\mathbf{R}\right| \\ - \Lambda'(\gamma - 1)\bar{w}(R_{LL}f/\bar{N} + R_{GL}g/\widehat{N})fR_Lf''\widehat{N}/\left|\mathbf{R}\right|.\end{aligned}$$

This is unambiguously positive iff $R_{GG}g/\widehat{N} + R_{GL}f/\bar{N} < 0$ and $R_{LL}f/\bar{N} + R_{GL}g/\widehat{N} < 0$. Both conditions apply to most revenue/production functions, and follow from assuming that effect of increasing the number of standard-time workers is going effect the marginal product of a standard-time worker more than it does an overtime worker, and vice versa.

Analogous expressions for the third channel are not ambiguous for no-scale-effects revenue functions. However, it is easy to show that for $R = \widehat{N}^{\beta_2}\bar{N}^{\beta_1}\widehat{H}^{\alpha_2}\bar{H}^{\alpha_1}$,

$$\frac{\partial \widehat{N}}{\partial \bar{H}}\frac{\bar{H}}{\widehat{N}} = \frac{\partial \bar{N}}{\partial \bar{H}}\frac{\bar{H}}{\bar{N}} = \frac{\alpha_1}{1 - \beta_1 - \beta_2} > 0 \qquad \frac{\partial p}{\partial \bar{H}} = 0.$$

All three make intuitive sense, and if it is the case that $\partial p/\partial \bar{H}$ is zero, or very small, for more general revenue functions then the overall prediction that a cut in standard hours lowers the proportion of overtime workers in a firm still holds up.

References

Andrews, M. & Simmons, R. (2001), 'Friday may never be the same again: some results on work sharing from union-firm bargaining models', *Scottish Journal of Political Economy* **48**, 488–516.

Arellano, M. & Honoré, B. (2001), Panel data models: Some recent developments, *in* J. Heckman & E. Leamer, eds, 'Handbook of Econometrics', Vol. 5, Elsevier, Amsterdam, chapter 53, pp. 3229–96.

Bauer, T. & Zimmermann, K. (1999), 'Overtime work and overtime compensation in Germany', *Scottish Journal of Political Economy* **46**, 419–36.

Bell, D. & Hart, R. (1999), Overtime in an unregulated labour market, Discussion Paper No. 44, Institute for the Study of Labor (IZA), June.

Blundell, R. & MaCurdy, T. (1999), Labor supply: a review of alternative approaches, *in* O. Ashenfelter & D. Card, eds, 'Handbook of Labor Economics', Vol. 3B, Elsevier, Amsterdam, chapter 27, pp. 1559–666.

Brunello, G. (1989), 'The employment effects of shorter working hours: an application to Japanese data', *Economica* **56**, 473–86.

Calmfors, L. & Hoel, M. (1988), 'Work sharing and overtime', *Scandinavian Journal of Economics* **90**, 45–62.

Crépon, B. & Kramarz, F. (2002), 'Employed 40 hours or not-employed 39: lessons from the 1982 mandatory reduction of the workweek', *Journal of Political Economy* **110**, 1355–89.

Deaton, A. (1997), *The Analysis of Household Surveys: a Microeconometric Approach to Development Policy*, John Hopkins University Press, Baltimore.

Ehrenberg, R. (1971), *Fringe Benefits and Overtime Behaviour*, Lexington Books, Lexington, Mass.

European Industrial Relations Review (2001), 'New 35 hour week for small companies', **335**, 26–27.

European Industrial Relations Review (2002*a*), 'Amendments to working time law in progress', **346**, 6.

European Industrial Relations Review (2002*b*), 'New law relaxes implementation of 35 hour week', **335**, 15–17.

European Industrial Relations Review (2003*a*), 'IG Metall ends 35 hour week campaign', **354**, 5–6.

European Industrial Relations Review (2003*b*), 'IG Metall pushes for 35 hour week in the East', **352**, 15–17.

Hamermesh, D. (1993), *Labor Demand*, Princeton University Press, Princeton.

Hart, R. (1984), *The Economics of Non-wage Labour Costs*, Allen and Unwin, London.

Hart, R. (1987), *Working Time and Employment*, Allen and Unwin, London, chapter 5.

Hart, R., Malley, J. & Ruffell, R. (1996), 'What shapes are overtime premium schedules? some evidence from Japan, the UK, and the US', *Economics Letters* **53**, 97–102.

Hart, R. & Ruffell, R. (1993), 'The cost of overtime hours in British production industries', *Economica* **60**, 183–202.

Hart, R. & Wilson, N. (1988), The demand for workers and hours: micro evidence for the UK metal working industry, *in* R. Hart, ed., 'Employment, Unemployment, and Labor Utilisation', Allen and Unwin, London, chapter 9.

Hernanz, V., Izquierdo, M. & Jimeno, J. (1999), Working hours and employment at the firm level: evidence from a panel of Spanish manufacturing firms, mimeo, Universidad de Alcalá and FEDEA, November.

Hübler, O. (1989), 'Individual overtime functions with double correction for selectivity bias', *Economics Letters* **29**, 87–90.

Hübler, O. & Meyer, W. (1997), 'Ueberstunden im verarbeitenden gewerbe niedersachsens', *Beitraege zur Arbeitsmarkt- und Berufsforschung* **204**, 227–254.

Hunt, J. (1998), 'Hours reductions as work-sharing', *Brookings Papers on Economic Activity* **1:1998**, 349–81.

Hunt, J. (1999), 'Has work-sharing worked in Germany?', *Quarterly Journal of Economics* **114**, 117–48.

Jacobson, T. & Ohlsson, H. (2000), 'Working time, employment, and work sharing: evidence fron Sweden', *Empirical Economics* **25**, 169–87.

Kalwij, A. & Gregory, M. (2000), Overtime hours in Great Britain over the period 1975–1999: a panel data analysis, Discussion Paper No. 153, Institute for the Study of Labor (IZA), May.

Kapteyn, A., Kalwij, A. & Zaidi, A. (2000), The myth of worksharing, Discussion Paper No. 188, Institute for the Study of Labor (IZA), August.

König, H. & Pohlmeier, W. (1989), 'Worksharing and factor prices: a comparison of three flexible functional forms for non-linear cost schemes', *Journal of Institutional and Theoretical Economics* **145**, 343–57.

Leslie, D. (1991), 'Modelling hours of work in a labour services function', *Scottish Journal of Political Economy* **38**, 19–31.

Schank, T. (2001), Estimating worksharing and related issues: evidence from German establishment panel data, PhD thesis, University of Manchester.

Schmidt-Sorensen, J. (1991), 'An efficiency-wage-hours model and shorter working hours', *Scottish Journal of Political Economy* **38**, 113–31.

The Guardian (2003), 'Labouring over the 35 hour week', **Saturday October 11**.

Wooldridge, J. (1999), *Introductory econometrics: an introductory approach*, Thomson.

Wooldridge, J. (2002), *Econometric analysis of cross section and panel data*, MIT Press.

Table 1: Standard hours elasticities on actual hours, employment, and probability of working overtime: microeconometric evidence

Study	Data	$\eta_{H\bar{H}}$	$\eta_{N\bar{H}}$	$\frac{\partial \Pr(V>0)}{\partial \bar{H}}$ [a]
Hart & Wilson (1988)	52 UK engineering firms, panel, 1978–82	0.80	0.41/−0.49[b]	
Bell & Hart (1999)	New Earnings Survey, 24029 males, 1996; GB	0.798[c]		−0.012
Own regressions	New Earnings Survey, 34657 manual males, 1978; GB	0.895[c]		−0.0053
Own regressions	New Earnings Survey, 31360 manual males, 1985; GB	1.056[c]		−0.0005
Kalwij & Gregory (2000)	New Earnings Survey, 1975–99; GB	0.98[d]		
Bauer & Zimmermann (1999)	German Socio-Economic Panel, 17332 individuals, 1984–97; West Germany	1.034[c]		0.0020
Hübler (1989)	German Socio-Economic Panel, 1031 individuals, 1984; West Germany	0.924[e]		≈ −0.0075
Hübler & Meyer (1997)	Panel of firms (1024 in 94; 849 in 95); West Germany	0.99		
Hunt (1999)	German Socio-Economic Panel, 4386 workers, mnfg/service sectors, 1984-94; West Germany	[0.70,0.85]		[−0.0029, −0.0017][f]
Hunt (1999)	30 manufacturing industries, bi-annual pooled, 1982–93; West Germany	0.90	–0.50[g]	
Hernanz et al. (1999)	Spanish EESE Panel, 7300 firms, manufacturing, 1990–97	1.09	–0.005	
Crépon & Kramarz (2002)	French Labour Force Survey, panel, 1977-87		[0.8,1.6][h]	

[a] Except for Bauer & Zimmermann (1999) and Hunt (1999), estimated as a Probit.

[b] $\eta_{N\bar{H}} = 0.41$ for firms who offer overtime, $\eta_{N\bar{H}} = -0.49$ otherwise.

[c] Is the regression parameter in a Tobit, ie estimates $\partial H^*/\partial \bar{H}$, where H^* is unconstrained choice of hours.

[d] ML Fixed Effects Tobit.

[e] As c, but uses Heckman-corrected truncated regression.

[f] Linear probability model with individual fixed-effects.

[g] Insignificant. Changes to 0.71 in a 10-industry panel, 1984–94.

[h] Effect of reducing $\bar{H}$ from 40 to 39 hours in 1982 on prob of employment to non-employment transition.

Table 2: Average weights (ω) for different plant sizes (N)

	West	East
$1 \leq N < 5$	1209.70	240.91
$5 \leq N < 10$	1065.38	178.00
$10 \leq N < 20$	591.88	95.20
$20 \leq N < 50$	234.62	38.50
$50 \leq N < 100$	112.92	19.44
$100 \leq N < 200$	56.44	9.09
$200 \leq N < 500$	22.35	4.01
$500 \leq N < 1000$	10.91	2.41
$1000 \leq N < 5000$	3.36	2.00
$5000 \leq N$	2.05	1.26
All	382.19	83.38

Table 3: Descriptive Statistics

	Weighted[a]											Unweighted	
	West						*East*						
	93	95	96	97	98	99	96	97	98	99	All	N-Reg.	p-Reg.
Standard-time plants ($D = 0$)													
N	11.7	10.9	11.5	11.8	11.7	10.9	11.9	12.3	12.0	10.4	11.4	127	105
$\bar{H}$	38.9	39.5	38.9	38.5	39.2	39.5	39.8	39.7	39.9	40.2	39.2	39.0	39.1
w	14.5	15.3	15.6	14.7	14.2	14.3	12.4	12.2	12.1	11.7	14.3	17.2	16.7
Y	5.77	6.02	5.63	9.17	6.08	6.70	3.71	3.53	3.99	3.70	6.17	114	96.8
I	.328	.253	.222	.108	.111	.115	.248	.219	.249	.181	.176	3.65	1.84
$B = 1, 2$	.699	.581	.580	.551	.500	.443	.451	.443	.333	.271	.529	.642	.614
$\sum \omega_i{}^d$												9109	7471
Overtime plants ($D = 1, 2$)													
N	36.5	41.3	38.8	37.7	37.1	42.6	26.6	25.1	26.5	29.2	37.3	589	393
p			.569	.617	.582	.509[c]	.643	.680	.608	.538[c]	.581	.414	.417
$\bar{H}$	39.5	39.0	39.6	39.3	39.4	39.0	40.5	40.0	40.2	40.1	39.4	38.3	38.5
V			3.31	2.69	2.90	2.19[c]	3.05	2.86	2.90	1.71[c]	2.97	2.82	2.77
H			40.6	40.2	40.8	39.9[c]	42.3	41.6	41.2	40.7[c]	40.8	39.3	39.5
w	19.6	20.6	19.6	19.8	19.4	20.7	15.0	14.9	15.1	16.0	19.3	24.8	23.6
Y	29.6	32.3	29.3	34.5	37.1	38.5	8.39	8.13	10.4	13.0	30.3	895	539
I	1.05	1.20	1.04	.660	.568	.695	.764	.665	.644	.602	.822	18.4	11.4
$B = 1, 2$	.768	.714	.637	.625	.556	.585	.495	.503	.411	.359	.625	.810	.768
$\% \, p_i = 1 \quad (D = 2 \,\vert\, D = 1, 2)$			.277	.312	.275	.198	.355	.387	.295	.217	.278	.074	.089
$\sum \omega_i{}^d$												9487	5692

to be continued…

	Weighted[a]											Unweighted	
	West						*East*						
	93	95	96	97	98	99	96	97	98	99	All	N-Reg.	p-Reg.
All plants													
N^b	19.2	18.9	18.6	18.1	18.1	17.9	16.2	15.4	15.1	14.1	18.0	363	230
p			.132	.129	.123	.098[c]	.178	.156	.114	.093[c]	.123	.167	.181
$\bar{H}$	39.1	39.3	39.1	38.7	39.2	39.4	40.0	39.7	40.0	40.1	39.3	38.6	38.8
w	16.1	16.7	16.7	16.0	15.6	15.8	13.1	12.9	12.7	12.6	15.7	21.1	19.7
Y	13.4	13.2	11.8	15.6	14.0	13.9	5.07	4.70	5.34	5.61	12.4	515	291
I	.579	.549	.434	.241	.225	.241	.396	.328	.333	.263	.342	11.2	6.00
B	.719	.616	.595	.567	.513	.475	.463	.467	.349	.288	.553	.728	.681
$D = 1,2$	.300	.261	.260	.247	.256	.221	.292	.248	.214	.194	.254	.510	.432
$\sum \omega_i{}^d$	1595	1625	1634	1639	1644	1653	391	398	401	427		18596[e]	13163[e]
$\sum \omega_i N_i{}^f$	30640	30674	30369	29732	29830	29576	6329	6140	6054	6004			

[a] The weighted columns refer to the whole sample but the unweighted columns to the two regression samples.
[b] Ratio of c to b. See Equation (12) of main text.
[c] In 1999, p refers to employees working paid overtime, V refers to paid overtime and H has been calculated accordingly.
[d] Total number of observations (in thousands for weighted figures). Estimates number of plants in population.
[e] Regression sample sizes. See Tables 5 and beyond.
[f] Total number of employees (in thousands). Estimates population employment.

Table 4: Observations per year and per plant

	N–reg.[a]	p–reg.[b]
No. of plants with 1 observation	0	0
No. of plants with 2 observations	1266	1152
No. of plants with 3 observations	995	1101
No. of plants with 4 observations	1763	1889
No. of plants with 5 observations	645	0
No. of plants with 6 observations	467	0
No. of plants	5136	4142
No. of observations in 1993	1208	0
No. of observations in 1995	2174	0
No. of observations in 1996	4321	3472
No. of observations in 1997	4219	3742
No. of observations in 1998	3673	3265
No. of observations in 1999	3001	2684
No. of plant-years	18596	13163
No. of differences	13315	9021
No of non-zero $\Delta\bar{h}$ obs	2851	1176

[a]Sample corresponds to regressions reported in Table 5.
[b]Sample corresponds to regressions reported in Table 7.

Table 5: Employment regressions in differences (1993-99, excluding 1994)[a]

	All		West				East			
			Ag.,Mnfg.		Services		Ag.,Mnfg.		Services	
	coeff	std error	coeff	std error	coeff	std error	coeff	std error	coeff	std error
BaseFD										
obs, clusters	13315	1044	4561	258	3823	416	2656	146	2275	227
FD, unweighted										
$\Delta\bar{h}$	−0.061	(0.073)	0.082	(0.140)	0.064	(0.104)	−0.754	(0.239)	−0.176	(0.150)
SC test[b]	0.010	(0.011)	0.085	(0.018)	−0.044	(0.019)	0.016	(0.027)	−0.062	(0.029)
FD, weighted										
$\Delta\bar{h}$	−0.023	(0.097)	−0.034	(0.254)	0.063	(0.118)	−0.914	(0.292)	−0.164	(0.165)
SC test	−0.065	(0.027)	0.033	(0.058)	−0.110	(0.030)	−0.053	(0.055)	−0.131	(0.068)
Hausman test of unweighted v weighted[c]										
F(1,.) [pv]	0.356	[0.551]	0.300	[0.584]	0.001	[0.982]	0.914	[0.341]	0.032	[0.859]

to be continued...

[a]Estimates of Equation (15). Other covariates included are bargaining dummy, a demand shock, investment and proportion of: female employees, part-time employees, skilled employees/apprentices, employees working on shifts, employees working on Saturdays, employees working on sundays, employees working on a flexible working time schedule. All variables are differenced except for investment. Standard errors in ( ).

[b]Wooldridge's Serial Correlation test: regress $\widehat{\Delta} u_{it}$ on $\widehat{\Delta} u_{i,t-1}$.

[c]Hausman test for whether weighted/unweighted effects of $\bar{h}$ differ. Distributed $F(1,.)$ under H_0. p-values in [ ].

	All		West				East			
			Ag.,Mnfg.		Services		Ag.,Mnfg.		Services	
	coeff	std error	coeff	std error	coeff	std error	coeff	std error	coeff	std error
Endogeneity test:[d]										
FD, unweighted										
$\Delta\bar{h}$	−0.059	(0.073)	0.083	(0.139)	0.063	(0.104)	−0.731	(0.238)	−0.163	(0.154)
$\bar{h}$	−0.014	(0.007)	−0.007	(0.015)	−0.005	(0.010)	−0.026	(0.012)	−0.028	(0.013)
IV is $1\{B=0,1\}\bar{h}+1\{B=2\}\bar{h}^*$										
obs, clusters	12547	1015	4484	256	3555	398	2530	143	1978	221
FDIV, unweighted										
$\Delta\bar{h}$	−0.033	(0.090)	−0.022	(0.235)	0.134	(0.117)	−0.767	(0.305)	−0.139	(0.163)
First stage FDIV										
$\Delta\bar{h}^*$	0.854	(0.020)	0.685	(0.042)	0.924	(0.020)	0.840	(0.027)	0.971	(0.009)
IV is $1\{B=1\}\bar{h}+1\{B=0,2\}\bar{h}^*$										
obs, clusters	10905	943	4364	254	2859	365	2282	142	1400	185
FDIV, unweighted										
$\Delta\bar{h}$	−0.044	(0.154)	−0.285	(0.265)	0.335	(0.220)	−0.839	(0.511)	0.042	(0.287)
First stage FDIV										
$\Delta\bar{h}^*$	0.627	(0.037)	0.585	(0.046)	0.626	(0.074)	0.679	(0.051)	0.869	(0.049)

[d]Wooldridge's endogeneity test. Signifiance of adding $\bar{h}_{it}$ to Equation (15).

Table 6: Differenced employment regressions by plant size and bargaining agreement (1993–99, excluding 1994)[a]

	All		West				East			
			Ag.,Mnfg.		Services		Ag.,Mnfg.		Services	
	coeff	std error	coeff	std error	coeff	std error	coeff	std error	coeff	std error
obs, clusters	13315	1044	4561	258	3823	416	2656	146	2275	227
FD, unweighted										
$\Delta[1\{N_{it} < 15\}\bar{h}_{it}]$	−0.167	(0.097)	−0.068	(0.333)	0.000	(0.123)	−1.072	(0.335)	−0.147	(0.173)
$\Delta[1\{15 \leq N_{it} < 100\}\bar{h}_{it}]$	0.065	(0.171)	0.212	(0.278)	0.213	(0.252)	−0.579	(0.455)	−0.135	(0.492)
$\Delta[1\{N_{it} \geq 100\}\bar{h}_{it}]$	0.268	(0.204)	0.158	(0.188)	−0.203	(0.510)	−0.007	(0.623)	0.599	(0.771)
$\Delta[1\{B_{it} = 1, 2\}\bar{h}_{it}]$	−0.040	(0.123)	−0.026	(0.139)	0.090	(0.146)	−0.071	(0.424)	−0.373	(0.519)
SC test	0.010	(0.011)	0.085	(0.018)	−0.043	(0.019)	0.017	(0.027)	−0.062	(0.029)
IV is $1\{B = 1\}\bar{h} + 1\{B = 0, 2\}\bar{h}^*$										
obs, clusters	10905	943	4364	254	2859	365	2282	142	1400	185
FDIV, unweighted										
$\Delta[1\{N_{it} < 15\}\bar{h}_{it}]$	−0.123	(0.124)	−0.277	(0.613)	0.074	(0.130)	−1.090	(0.422)	−0.108	(0.207)
$\Delta[1\{15 \leq N_{it} < 100\}\bar{h}_{it}]$	0.187	(0.194)	0.093	(0.339)	0.420	(0.282)	−0.537	(0.604)	0.034	(0.552)
$\Delta[1\{N_{it} \geq 100\}\bar{h}_{it}]$	0.335	(0.259)	0.097	(0.287)	0.068	(0.661)	−0.569	(0.782)	1.045	(1.071)
$\Delta[1\{B_{it} = 1, 2\}\bar{h}_{it}]$	−0.123	(0.137)	0.017	(0.188)	−0.038	(0.147)	0.161	(0.530)	−0.659	(0.646)
No of non-zero $\Delta\bar{h}$ obs										
$1\{N_{it} < 15\}$	1075		281		506		116		152	
$1\{15 \leq N_{it} < 100\}$	796		265		258		177		96	
$1\{N_{it} \geq 100\}$	980		630		142		141		67	
All	2851		1176		906		434		335	

[a]Estimates of Equation (15), with interactions between standard, plant-size, and bargaining dummy. For FDIV, each $\bar{h}$ variable is instrumented by the corresponding h^* variable.

Table 7: Regressions for total employment and proportion of overtime workers (1996–99), unweighted

	All		West				East			
			Ag.,Mnfg.		Services		Ag.,Mnfg.		Services	
	coeff	std error	coeff	std error	coeff	std error	coeff	std error	coeff	std error
Employment, differences[a]										
Dependent variable: Δn_{it}										
$\Delta \bar{h}_{it}$	−.121	(.086)	−.016	(.204)	.156	(.133)	−.780	(.240)	−.186	(.156)
No of plants		4142		1079		1023		1089		951
No of plant-years		13163		3364		3190		3551		3058
No of differences		9021		2285		2167		2462		2107
No of non-zero $\Delta \bar{h}$ obs		1660		416		515		408		321
Proportion of overtime workers, two-sided Tobit[b]										
Dependent variable: p_{it}										
$\bar{h}_{it}$	.345	(.167)	−.045	(.227)	.710	(.271)	.189	(.404)	.678	(.593)
$\partial p/\partial \bar{h}$		.134		−.024		.226		.080		.177
$\partial Pr(D=0)/\partial \bar{h}$		−.236		.038		−.459		−.125		−.327
$\partial Pr(D=2)/\partial \bar{h}$		.034		−.005		.043		.027		.053
No of plant-years		13163		3364		3190		3551		3058
Plant-years, $D=0$		7480		1438		2047		1874		2121
Plant-years, $D=1$		5179		1811		1045		1514		809
Plant-years, $D=2$		504		115		98		163		128

to be continued…

[a] Estimates of Equation (15). Standard errors in ( ). Note that the dependent variable is $\Delta \log N_{it}$ rather than ΔN_{it}.
[b] Estimates of Equation (17). Same covariates as in employment equation.

	All		West				East			
			Ag.,Mnfg.		Services		Ag.,Mnfg.		Services	
	coeff	std error	coeff	std error	coeff	std error	coeff	std error	coeff	std error
Standard hours elasticities[c]										
Total employment, $\eta_{N\bar{H}}$		$-.121$		$-.016$		$-.156$		$-.780$		$-.186$
Standard time workers, $\eta_{\bar{N}\bar{H}}$		$-.238$		.007		$-.049$		$-.845$		$-.323$
Overtime workers, $\eta_{\hat{N}\bar{H}}$		.199		$-.076$		.739		$-.605$		.230
(Sample mean of p)		(.181)		(.227)		(.139)		(.216)		(.130)
(Sample mean of $p \mid D = 1, 2$)		(.418)		(.397)		(.388)		(.458)		(.426)
(Sample mean of $p \mid D = 0, 1$)		(.148)		(.200)		(.112)		(.179)		(.092)

[c] See Equations (21) and (22) of main text.

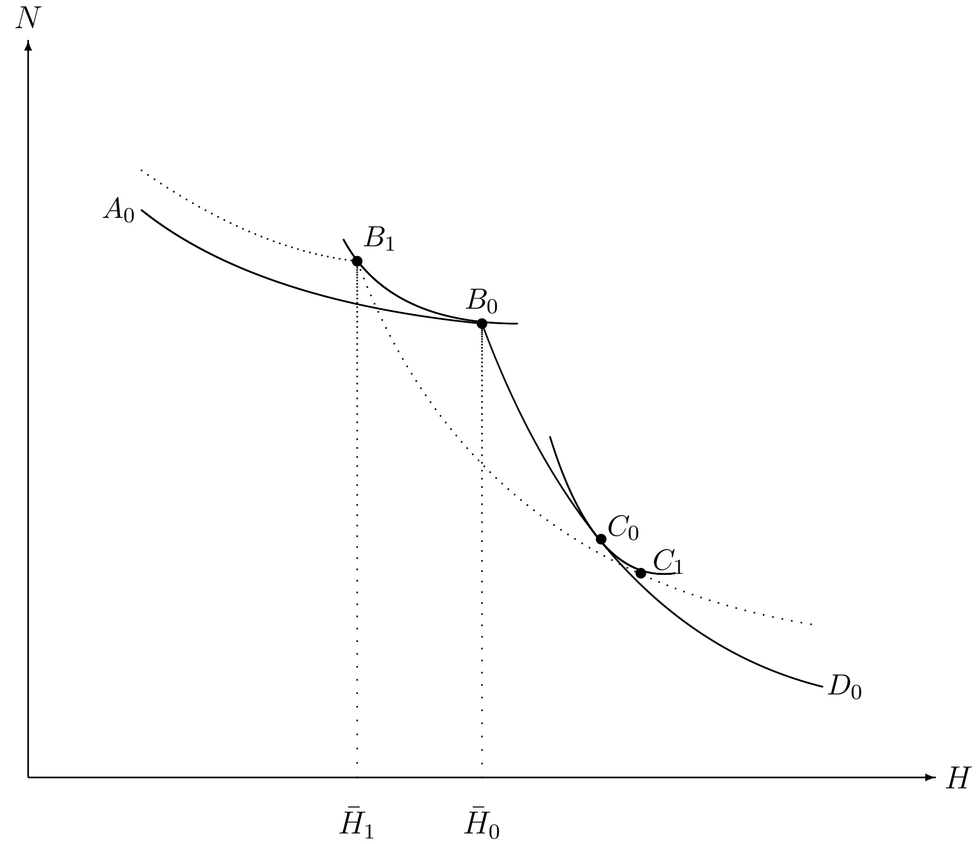

Figure 1: Possible solutions and substitution effects of a cut in the workweek

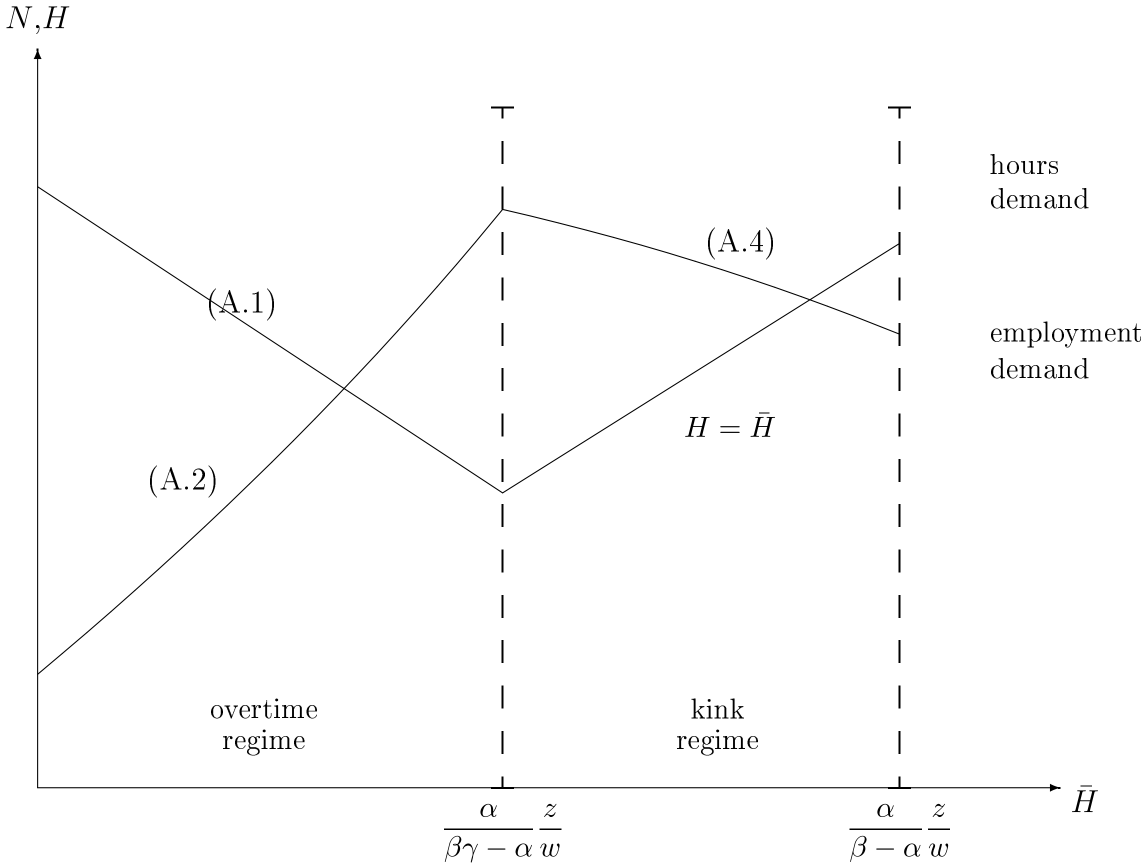

Figure 2: Employment and hours functions (Cobb-Douglas example)

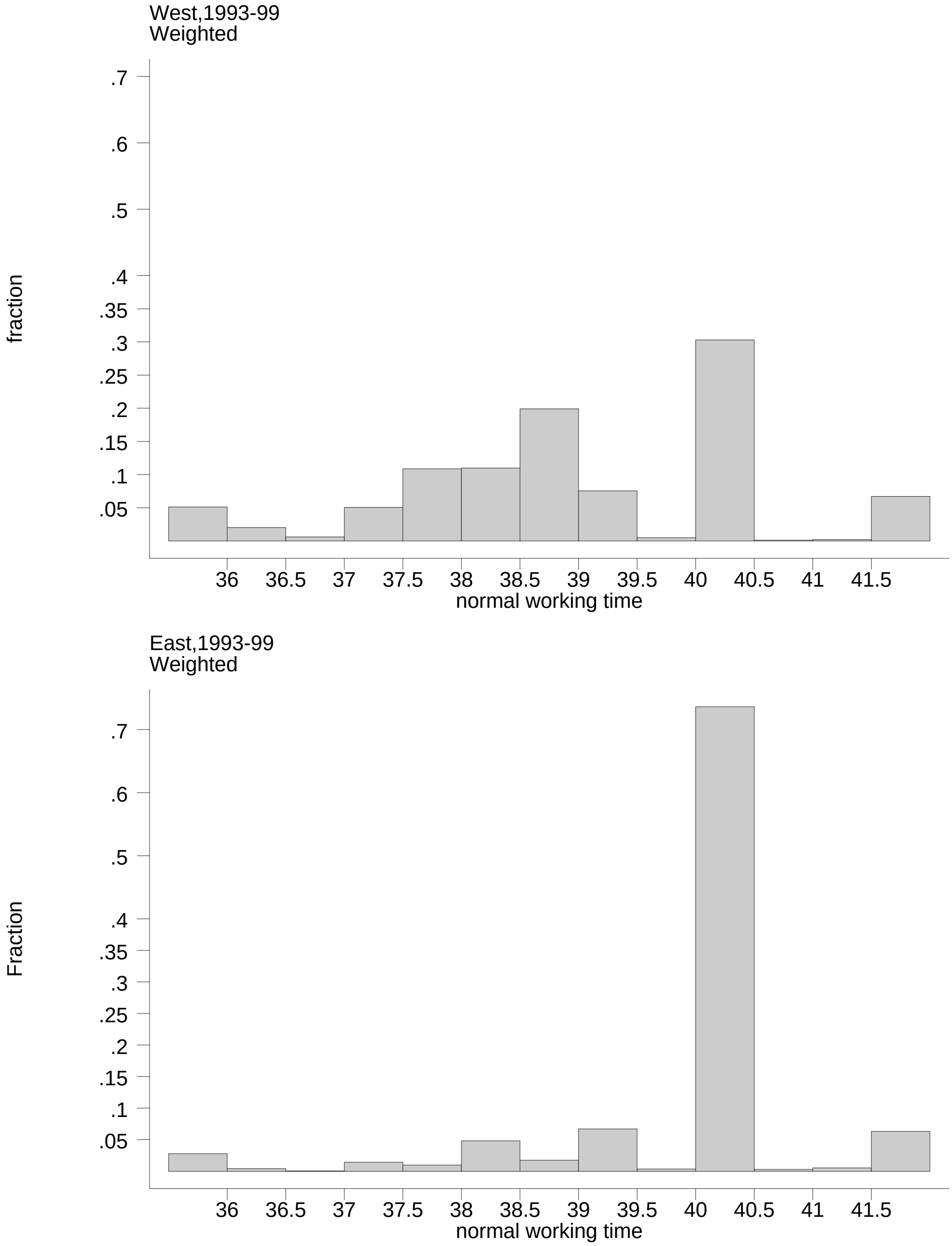

Figure 3: Weighted distribution of normal working time

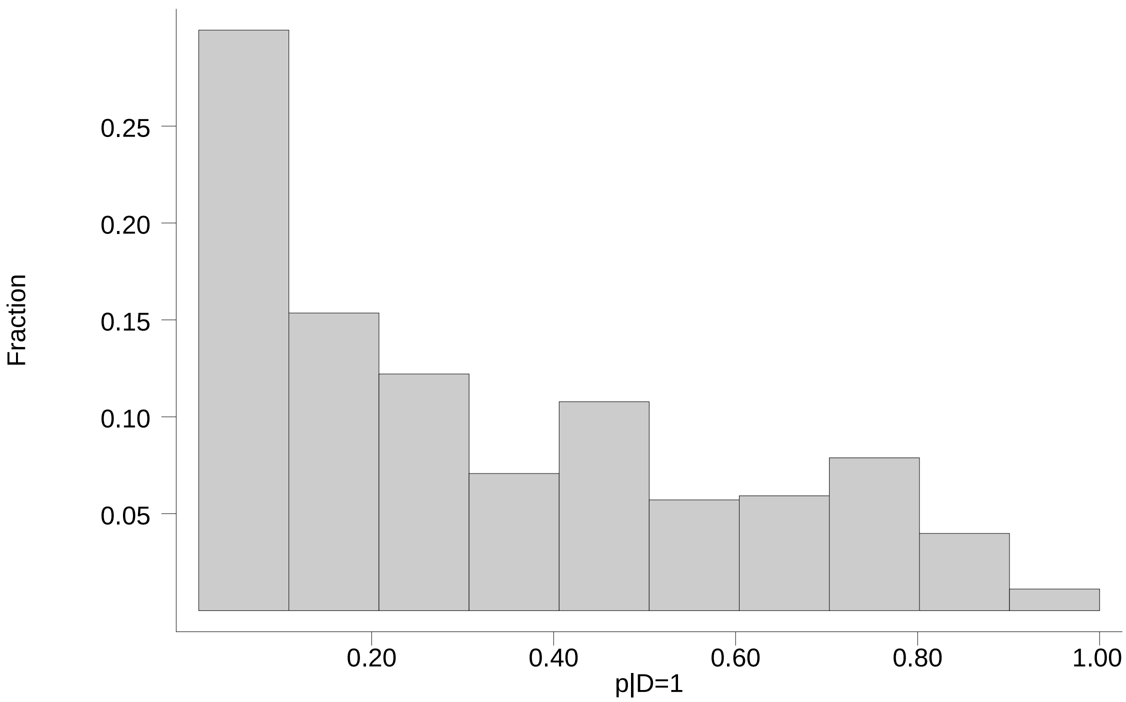

Figure 4: Truncated distribution of proportion of overtime workers in a plant

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