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Who benefits from firm success? Heterogeneous rent-sharing in New Zealand

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Abstract

We continue our examination of inclusive growth at the firm level by examining heterogeneity in rent sharing in New Zealand using linked employer-employee data. We test for heterogeneity in rent sharing across a range of worker and firm characteristics including gender, ethnicity, age, qualifications, tenure, firm size, firm age, and industry. We also refine our measure of quasi-rents and estimate the level of excess quasi-rents per worker, or the amount of rents above the threshold beyond which rent sharing occurs. We find that between 20% and 30% of workers are in firms that earn zero excess rents. These workers are concentrated in the hospitality, administrative services, and retail industries and are more likely to be women, to be Māori or Pacific peoples, and have lower-level qualifications. We find an overall rent-sharing elasticity of 0.03, which is equivalent to a $38 increase in annual wages in response to a $1,000 increase in excess rents per worker. We find differences in rent sharing by levels of highest qualification, tenure, and ethnicity. We find no differences in rent sharing by firm size or firm age. Rent sharing is similar across industries, with workers in most industries receiving between $1,500 and $2,000 of rents per year. The auxiliary finance and professional, scientific, and technical services sectors share the most, while grocery retailing, food and beverage manufacturing and utilities share the least. Insurance type behaviour by firms is consistent with the variation in rent sharing across industries, although differences in bargaining power are also likely to play a role in explaining differences in rent sharing across groups.

JEL codes

J31 Wage Level and Structure - Wage Differentials; J71 Discrimination; J10 Demographic economics – General; D22 Firm Behavior: Empirical Analysis.

Keywords

Wage determination; Rent-sharing; imperfect competition.

Contents

1 Introduction 1

2 Background 3

3 Data and general patterns 7

3.1 Data 7

3.2 General patterns 12

4 Estimation approach 15

5 Results 19

5.1 How are the rents shared with different workers? 20

5.2 Which types of firms share more rents? 22

6 Conclusion and discussion 26

References 29

List of Tables

Table 1: Summary statistics for estimation sample – jobs in firms with positive excess rents 39

Table 2: Average wages and rents by worker/job characteristics 41

Table 3: The relationship between rents and wages - moving from the firm-level to the job-level 43

Table 4: Heterogeneous rent-sharing - how rents are shared with different workers 44

Table 5: Heterogeneous rent-sharing - gender-specific regressions 46

Table 6: Heterogenous rent sharing by firm characteristics 48

List of Figures

Figure 1: The piecewise linear relationship between quasi-rents per worker and firm wage premiums, 2018 32

Figure 2: Percent of firms and employment that earn zero excess quasi-rents 33

Figure 3: Excess rents per worker and wages by industry 34

Figure 4: Rent-sharing elasticities by industry 35

Figure 5: Rent sharing per dollar by industry 36

Figure 6: Average amount of rents received per worker per year by industry 37

Figure 7: Rent sharing and firm-level volatility in rents per worker 38

1 Introduction

The issue of power dynamics in the labour market is attracting increasing attention from both policy makers and researchers throughout the developed world (e.g., Blanchflower 2019; Stansbury & Summers 2020; Manning 2011). Across almost all OECD countries, the labour income share has fallen, recent wage growth has been low relative to historic experience, and union membership is in decline. Several factors seem to be serving to undermine the power of workers in the labour market, from increasing automation, offshoring, the rise of the gig economy and other forms of non-standard work.

A range of recent policy changes in New Zealand can be thought of as addressing low bargaining power and raising wages of historically disadvantaged groups of workers. These include removing 90-day trial periods for firms with more than 20 employees, further increases in the minimum wage, a range of pay equity settlements (the first and largest of which was for aged care and support workers), and the recently announced Fair Pay Agreements.

In this paper, we extend Allan & Maré (2021) and continue our examination of inclusive growth at the firm level in New Zealand by examining the extent of heterogeneous rent sharing within firms. We look at whether rent sharing differs across workers with different characteristics and whether different types of firms are more or less likely to share any rents. This provides further insights into the drivers of wage gaps between different groups (e.g., men and women, Māori and NZ European) by looking for the presence of these gaps within firms. This work builds on our previous work looking at overall patterns in the pass-through of firm performance to wages both across and within firms, as well as the role of worker sorting and rent sharing in explaining overall pass-through (Allan & Maré, 2021).

Our previous work shows a positive relationship between a measure of quasi-rents per worker and average wages at the firm level, with average annual wages increasing by approximately $80 in response to a $1,000 increase in quasi-rents per worker at the same firm. In this work, we consider heterogeneity in rent sharing across several worker and firm characteristics, asking which workers benefit from improvements in firm performance.

We find that approximately 23% of workers are in firms where there are no rents to share. These workers are concentrated in the hospitality, administrative services, and retail industries and are more likely to be women, to be Māori or Pacific peoples, are younger and have lower levels of educational attainment.

We further refine the analysis from Allan & Maré (2021). Our baseline rent sharing estimate, for firms with rents to share, shows that workers receive an average of $38 for a $1,000 increase in excess rents per worker. This is lower than our previous estimate but, given the relative imprecision of our previous estimate, within the 95% confidence interval.

We find that workers with higher qualifications and longer tenure have higher rent-sharing elasticities than workers with lower qualifications or shorter tenure. Conversely, Māori and Pacific workers, and Pacific women in particular, do not benefit from rent sharing to the same extent as European workers.

Our results suggest that insurance behaviour may be a contributing factor behind differences in rent-sharing elasticities across groups. However, the association between rent-sharing elasticities and measures of volatility in firm performance is relatively weak, suggesting other explanations are also important. These factors could include differences in bargaining power across groups related to the relative supply and demand of different types of workers, differences in the extent of monopsony power that firms hold over different workers, differences in attitudes, experience and skills in wage bargaining, differences in attitudes to pay transparency, or discrimination. Further research is needed to quantitatively assess the importance of these drivers in explaining differences in rent sharing across groups.

The rest of this paper is structured as follows. Section 2 presents some background on rent sharing and wage bargaining. Section 3 discusses the data we use in this work and presents some general patterns. Section 4 presents our empirical strategy, and results are presented in section 5. Section 6 concludes.

2 Background

The starting point for rent sharing studies is the observed correlation between firm performance and wages. An extensive literature exists that considers the relationship between firm performance and workers' wages.[1] These studies typically use bargaining models or monopsony models to establish a link between firm performance and wages. These are often motivated by concerns about increasing inequality (e.g., Barth et al. 2016), examining insider-outsider dynamics (e.g., Blanchflower et al. 1990), or explaining wage differentials across different groups of workers (e.g., Card et al. 2016; Sin et al. 2020).

A common theoretical framework to explain this correlation is wage bargaining, where firms and workers (or their representatives) bargain over wages and profits (e.g., Blanchflower et al. 1990; Abowd & Lemieux 1993; Card et al. 2014; Bell et al. 2019). A crucial component of a bargaining model is the existence of some surplus (profits or rents) to be bargained over. A stylised bargaining model can be described by an asymmetric Nash bargain (Manning 2011):

$$(p - w)^{1-\alpha}(w - b)^{\alpha} \tag{1}$$

Where p is the value of what the workers produce, w is the wage, b is the reservation or alternate wage, and α describes the relative bargaining power of the worker(s). In perfectly competitive output and labour markets, $p = w = b$, there is no surplus to share, and workers earn the reservation or alternate wage of the marginal worker, which is equal to the marginal revenue product. Firms want to maximise the amount of surplus they can take as profits ($p - w$), while workers want to maximise their wage and the gap between their wage and the reservation wage. The wage equation generated from this type of model is:

$$w = \alpha p + (1 - \alpha)b \tag{2}$$

where the wage depends on the value of what workers produce (p), the reservation wage (b), and the bargaining power of workers (α).

Another explanation for the observed correlation between firm performance and wages is worker sorting. If highly skilled (and therefore high wage) workers are more likely to work at better performing firms, this will generate a positive correlation between firm performance and wages. Studies of worker sorting, based on two-way fixed effect models introduced by Abowd et al. (1999), generally find a positive correlation between the worker fixed effect (a measure of worker quality) and firm fixed effects (a measure of firm wage premiums that are correlated with firm performance) (e.g., Abowd et al. 1999; Card et al. 2013, 2018; Maré & Hyslop 2006; Song et al. 2019; OECD 2021). Studies examining the role of firms in wage inequality highlight increased worker sorting as an important driver of increased inequality, both the movement of high wage workers to high wage firms (worker to firm sorting) and the movement of high-wage workers to firms that employ other high-wage workers

[1] See Card et al. (2018) for a summary.

(worker to worker sorting). However, OECD (2021) find that 35% of the *change* in between-firm wage inequality can be explained by worker sorting, showing that worker sorting is not the only explanation of the correlation between wages and firm performance. Differences in wage setting practices across firms, potentially related to the presence of quasi-rents and how they are distributed, remains as a candidate explanation.

The presence of a surplus implies firms have some form of market power. This could be in the form of product market power (i.e., firms face a downward sloping demand curve and set prices above marginal cost) or labour market power (i.e., firms face upward sloping labour supply curves and can set wages below the marginal product). The exercise of these forms of market power generates economic rents (profits in excess of the minimum required to remain in business) and these rents are the surplus that is available to be bargained over. Some studies of rent sharing use changes in the extent of market power to identify plausibly exogenous variation in rents. Kline et al. (2019) use information on patenting activity, Van Reenen (1996) use information on innovation, and Abowd & Lemieux (1993) use changes in import and export prices.

The ability to generate rents may be maintained by pervasive barriers to entry of competing firms. Rents may also be temporary, or 'quasi-rents', in that they will be eroded over time by the entry of competitors or through the removal of barriers to market entry (e.g., loss of intellectual property protections). Quasi-rents likely describe the situation for many firms that operate in markets with relatively low barriers to entry.

The temporary nature of a proportion of firm rents suggests that not all the rents in a particular year may be available to be shared. If firms are uncertain about whether good performance will continue, they may bank rents earned in good years to help them meet costs in bad years. This means firms can insulate workers from shocks to firm revenue by keeping wages relatively stable, providing a form of income insurance. In this literature, firm rents are typically modelled as a random walk type process with a permanent component and a transitory component (e.g., Guiso et al. 2005; Cardoso & Portela 2009; Juhn et al. 2018). Firms can then (partially) insulate their workers from both temporary and permanent shocks to performance by keeping wages steady and banking positive shocks to provide a buffer against negative shocks in the future. The empirical literature finds that firms provide strong insurance against temporary shocks to firm performance but only partial insurance against permanent shocks (Guiso et al. 2005; Cardoso & Portela 2009; Juhn et al. 2018). Cardoso & Portela (2009) find that managers receive less insurance against shocks to firm performance than workers. Similarly, Juhn et al. (2018) find that workers at the top of the firm earnings distribution receive less insurance.[2] Arai & Heyman (2009) find that the positive relationship between firm profits and wages exists only for firms with increasing profits, suggesting firms keep wages steady when profits are declining. The results from the insurance literature suggest that permanent rents are more likely to be shared. For some firms, temporary rents may represent most of the rents in any given year.

[2] Less insurance can be interpreted as greater rent sharing.

The final consideration for thinking about rent sharing is what determines relative bargaining power. Standard determinants include the presence of unions who bargain on workers' behalf. Bargaining power could also reflect a scarcity of workers with particular skills or attributes, relative to the demand for these workers. This gives some workers greater bargaining power in the labour market. Models of monopsony based on search frictions (dynamic monopsony) also suggest that incumbent workers have greater bargaining power when recruitment costs for firms are high relative to job search costs for workers (Manning, 2003). This suggests that bargaining power has a cyclical component – when unemployment is high and there are many workers looking for work, it is comparatively easy for firms to replace workers who leave or to expand, meaning workers will have lower bargaining power. It also suggests that where staff turnover is particularly costly to the firm (hiring costs, training costs, loss of specific human capital), incumbent workers have greater ability to extract rents as firms prefer to avoid these costs associated with turnover.

Search frictions are not the only source of monopsony power. Firms and workers differ along a number of dimensions and workers view jobs at different firms as imperfect substitutes i.e., workers have heterogenous preferences over the firms at which they would like to work (Card et al. (2018) is one example of a preference-based model of monopsony). This gives firms monopsony power as workers will be more reluctant to leave a firm for which they have a stronger preference. A further source of monopsony power is that workers may lack accurate knowledge about outside job options. Jäger et al. (2021) find that workers anchor their beliefs about wages in other firms at their current wage, meaning workers in low-wage firms are systematically underestimating wages elsewhere. This gives firms monopsony power as workers are unlikely to leave for higher paying jobs as they are not aware of them, giving the firm the ability to keep wages low.

Behavioural factors can also influence relative bargaining power. Several studies have shown differences in bargaining behaviour between men and women (e.g., willingness to bargain, starting points, accepting lower offers, Niederle & Vesterlund (2011) summarise findings from this literature). Cultural differences in attitudes to hierarchy and authority, or in experience with the dominant workplace culture, can also affect people's willingness to raise issues with those above them in the hierarchy, including asking for higher wages (The Auckland Co-Design Lab 2016; Equal Employment Opportunities Trust 2011). Differences in risk preferences could also play a role, where the need to provide sufficient and stable income for the household outweighs the perceived risk of 'rocking the boat' (e.g., The Auckland Co-Design Lab 2016).[3] A lack of pay transparency between workers can also reduce a worker's bargaining power. Ignorance of what similar workers earn makes it difficult to negotiate for higher wages.[4]

[3] Shigeoka (2019) finds that the macroeconomic conditions during youth have a long-lasting effect on risk preferences.

[4] Recent studies by Baker et al. (2019) for Canada and Bennedsen et al. (2020) for Denmark suggest pay transparency laws are useful in reducing the gender pay gap, while Gulyas et al. (2021) finds no evidence that pay transparency rules reduced the gender gap in Austria. Frey

Discrimination is also likely to play a role for certain groups and this may affect both the level of the wage but also the prospects for wage increases (e.g., from internal promotions) (e.g., Lang & Spitzer, 2020; Sin et al. 2020).

Card et al. (2018) provide an overview of the rent sharing literature and find that typical rent-sharing elasticities are between 0.05 and 0.15 i.e., a 10% increase in rents per worker is associated with a 0.5% to 1.5% increase in wages. Several studies have examined differential rent sharing. Card et al. (2016) use a combination of a rent sharing model and two-way fixed effects models to study the role of bargaining and sorting in explaining the gender wage gap in Portugal. They find that women are less likely to work in high-rent firms and that women tend to receive a smaller share of firm-specific rents than men. Sin et al. (2020) apply the methodology of Card et al. (2016) to New Zealand and find similar results. Card et al. (2018) take a related approach, comparing rent-sharing elasticities for less educated and more educated male workers in Portugal and find the rent-sharing elasticities are similar across the two groups. A recent OECD cross-country study finds large differences in rent sharing between low and high skilled workers in several countries (OECD 2021). They also find a gender gap in rent sharing, although this gap is smaller than the difference between low and high skilled workers.[5]

Some more targeted studies use exogenous, permanent shocks to available rents and estimate the wage response to these shocks. These studies typically find larger rent-sharing elasticities than those surveyed in Card et al. (2018). Kline et al. (2019) use firm patenting activity, along with an *ex-ante* assessment of the 'value' of a patent to generate exogenous variation in rents. They find that workers capture on average 30 cents of every dollar of patent-induced rents. This is concentrated among men, and workers in the top half of the firm earnings distribution. Other studies use changes in corporate tax rates as an exogenous change in the amount of available surplus (e.g., Fuest et al. 2018; Suárez et al. 2016). These studies find that workers bear between 20% and 50% of the tax incidence and that this incidence disproportionately falls on women and lower wage workers.

(2021) recommends that countries implement pay transparency laws and offers advice on their design and implementation.

[5] OECD (2021) report estimates for Costa Rica, Germany, Finland, France, Hungary, Japan, the Netherlands, and Portugal.

3 Data and general patterns

Our analysis requires information on firm performance, reliable earnings information of individual employees, and employee demographics. We construct two datasets, one with firm-year observations on firm performance, and one with job-year observations (individual at a firm in a year) with information on earnings and demographics. We link our firm-level data with our job-level data to create a final dataset with information on individual employee earnings, employee demographics, and firm performance.

3.1 Data

We use the same firm data as in Allan & Maré (2021), which is drawn from StatsNZ's Longitudinal Business Database (LBD) and Integrated Data Infrastructure (IDI).[6] The population of interest consists of all employees and employing private-for-profit firms in the measured sector. Our sample of firms is drawn from the LBD productivity tables, which contains annual firm-level information on gross output, intermediate expenditure, and the cost of capital (Fabling & Maré 2015a; 2019).[7] We have information on 319,299 firms over the period 2002-2018 for a total of 1,713,759 firm-year observations. As we are interested in exploring within-firm heterogeneity, we restrict attention to firms with at least 5 full-time equivalent (FTE) employees with usable wage information (defined in section 3.1.1). Firms with fewer than 5 workers will have less cross-sectional variation in the types of workers they employ and provide less information to identify these differences.

Employees are identified through the IDI labour tables (Fabling & Maré, 2015b). These tables contain monthly job-level information on all paid employees in New Zealand and are derived from monthly PAYE tax returns. We take information on monthly earnings, estimated full-time equivalent (FTE) labour input, age, gender, and the estimated components from a 2-way fixed effect model, similar to Abowd et al. (1999). We calculate tenure directly from the tax data, counting the number of months a worker has been with the firm during a particular employment spell.

We supplement our core data with additional information on the demographic characteristics of workers. We use information on ethnicity from the IDI personal details table. We take information on highest qualification from the 2013 and 2018 Censuses. Information on age and gender is available in the labour tables and is sourced from the IDI personal details table.[8]

[6] See Fabling & Sanderson (2016) for more information on the LBD.
[7] Firms are identified using the permanent enterprise number (PENT) of Fabling (2011). The productivity tables cover approximately 70% of the private-for-profit firm population (Fabling and Maré 2019).
[8] Our data come from the December 2019 archive of the LBD (ibuldd_clean_dec_2019) and the January 2020 refresh of the IDI (IDI_clean_20200120).

We restrict attention to workers between the ages of 25 and 70. Workers in this age range will tend to have stronger labour market attachment and be more likely to work full time. Most of these workers will have finished their main education spell and have fully entered the labour market. We further exclude observations in the top 1% of the value-added per-worker distribution as the values for these firms seem implausibly high. Our job-level sample contains 10,084,600 job-year observations on 1,751,800 individuals at 46,917 firms over the period 2002-2018.[9]

3.1.1 Measuring wages

We use annual full-time equivalent (FTE) earnings as our wage measure, as in Allan & Maré (2021). We build this measure from monthly job-level information and place several restrictions on the job-months that we use to calculate an individual's annual earnings. We exclude the first and last months of an individual's employment at a firm. Monthly earnings in the first and last month of employment are not necessarily a good proxy for the underlying wage rate given that many workers will work for a part of their first or last reporting month. There may also be payments associated with starting or ending a job which further reduce the usefulness of earnings in the first and last months of employment (e.g., signing bonus, pay out of accrued annual leave).

We also exclude job-months where the individual is obviously part time from the wage calculation. 'Obviously part-time' is defined as a worker that is earning less than they would earn working 40 hours per week at the statutory minimum wage. For these months it is unclear whether variation across individuals (or within individuals over time) are due to differences in wage rates or differences in hours worked. Including these job-months would lead us to overstate the variation in wages both across individuals and over time. In total, we exclude around 35% of all job-months, which account for 10%-12% of total wages (see Allan & Maré 2021).[10]

One difficulty in working with monthly earnings information without hours information is that highly paid part-time workers look like low paid full-time workers. Some highly paid part-time workers will not be excluded by the `obviously part-time' criterion as their monthly earnings will exceed full-time minimum wage earnings. This is particularly an issue when comparing wages across groups with different average hours worked (e.g., men and women). We have no information to identify highly paid part-time workers so they are included but we will underestimate their underlying wage rate.

We also need a proxy for a worker's 'reservation wage', or what a worker could expect to earn in the absence of any rent sharing. Ideally, this measure should reflect differences in worker quality, where highly skilled workers have a higher expected reservation wage, reflecting their higher productivity.

[9] The main reason behind the reduction in the number of firms covered is the firm-size restriction we place on our job-level dataset. In February 2020, nearly 90% of firms had 5 or fewer employees. See https://www.stats.govt.nz/information-releases/new-zealand-business-demography-statistics-at-february-2020

[10] The exclusion of part-time workers is the main driver behind the exclusion of job-months.

To get a proxy for a worker's reservation wage, we utilise the estimated components from a two-way fixed effect model for wages:

$$\ln w_{ijt} = a_i + Z_{ijt}\beta + \lambda_t + \varphi_j + \varepsilon_{ijt} \qquad (3)$$

Where w_{ijt} is the wage paid to worker i employed at firm j in year t, $Z_{ijt}\beta$ is a gender-specific quartic in age, a_i is a worker fixed-effect which measures the transferable worker wage premium, φ_j is a firm-fixed effect which captures firm-specific pay premia, λ_t are year fixed effects and ε_{ijt} is the error term. This model is estimated on all job-years.

The two-way fixed effect model allows us to separate wages into an individual component $a_i + Z_{ijt}\beta$, and a firm component $\varphi_j + \bar{\varepsilon}_{jt}$.[11] The individual component reflects worker quality (e.g. experience, skills) and measures the amount a worker could expect to be paid at a firm that pays a zero premium.[12] We use this as the basis of our measure of the reservation wage. The firm component captures differences in average wages across firms that are unrelated to the composition of its workforce. This might include differences in pay practices (e.g., minimum wage employer, efficiency wages, commission or bonuses) or differences in firm performance (i.e., rent sharing). We call the firm component the firm wage premium.

3.1.2 Measuring firm performance

Our key independent variable is a measure of firm performance. Allan & Maré (2021) used both value added per worker and quasi-rents per worker as performance measures. Value added measures the amount left over after paying for intermediate inputs but does not account for the cost of employing capital or labour. Quasi-rents account for both the cost of capital and employing workers at their reservation wages.

In this paper we focus on our measure of quasi-rents and extend this using the concept of 'excess' rents from Card et al. (2016) and used in Sin et al. (2020). Allan and Maré (2021) calculate quasi-rents per worker as:

$$QR = \frac{VA - (r + \delta)K - bL}{L} \qquad (4)$$

where VA is value added (gross output less intermediate consumption), $(r + \delta)K$ is the cost of capital, and bL is the reservation wage bill. Our estimate for $(r + \delta)K$ is taken

[11] $\bar{\varepsilon}_{jt}$ is the firm-level average of the residuals obtained from estimating equation 3.

[12] As shown by Abowd et al. (1999; 2002), only *relative* firm premiums are identified in these types of models. Therefore, the normalisation of the firm fixed effects is important. The estimates available in the labour tables are mean-zero i.e. the firm premium is measured relative to the average firm. Our measure of the reservation wage is the minimum a worker could expect to earn, so mean-zero firm fixed effects are inappropriate for our use here. We re-centre the distribution of the firm fixed effects such that the $\varphi_j = 0$ at the first percentile (see Allan & Maré (2021) for more details).

directly from the Fabling-Maré productivity tables.[13] We calculate the reservation wage bill using the individual-specific components from the two-way fixed effect model in equation 3. Specifically, we calculate average b for each firm-year as:

$$\bar{b}_{jt} = \bar{a}_{jt} + \bar{X}_{jt}\beta + \lambda_t \tag{5}$$

Where bars denote firm averages. $\bar{a}_{jt}$ is the (FTE weighted) average of the individual fixed effects across workers at the firm in year t, and $\bar{X}_{jt}\beta$ is the (FTE weighted) average covariate index across workers at the firm, and λ_t is the year effect, which captures the effect of macro conditions on wages. We multiply this average by total firm FTE employment to calculate the reservation wage bill.

We then follow Card et al. (2016) and define excess (log) rents as:

$$\ln EQR = \max\{0, \ln QR - \tau\} \tag{6}$$

where τ is some threshold level of rents below which they are not shared with workers as higher (lower) wages. This specification is motivated by an observed non-linear (kinked) relationship between rents and wage premiums whereby in low-QR firms, wages are unrelated to levels of QR but in firms with QR above the threshold, there is a positive relationship. Firms below this threshold are referred to as 'zero excess rent' firms.[14]

We calculate a measure of excess rents using a year-specific threshold, estimated from data on all firms with at least 5 FTE employees. Following the approach of Card et al. (2016), we group firms into 100 bins according to the level of QR, with each bin containing 1% of FTE employment. Bins are defined separately by year. Within each bin, we calculate the mean of $\ln QR$ and the mean (FTE-weighted) firm fixed effect. Figure 1 plots mean (log) quasi-rents per worker against mean firm fixed effects, in 100 percentile bins of quasi-rents per worker, for 2018. A kink is clear in the data around 9.8. Below this level, higher quasi-rents are not necessarily associated with higher wage premiums. Above this threshold, there is a clear linear relationship between quasi-rents per worker and wage premiums.

To get a more precise estimate of τ, we estimate the following model separately by year:

$$\varphi_p = a + \gamma * \max\{0, \ln QR_p - \ln \tau\} + e_p \tag{7}$$

where φ_p is the average firm wage premium for bin p, a is the intercept, which will equal the average wage premium in zero excess-rent firms, γ is the sharing parameter,

[13] The measure of capital included in the Fabling-Maré productivity tables measures the flow of capital services and is the sum of reported depreciation, rental payments for rented capital goods, and a proxy for borrowing costs. This proxy is 10% of the value of fixed assets.
[14] The wage setting process in these firms likely reflects institutional characteristics, such as statutory minimum wages.

τ is the threshold level of rents, and e_p the error term. The results from estimating equation 7 by non-linear least squares is shown in Table A 1 in Appendix A.[15]

The estimated threshold $(\ln \tau)$ for 2018 is 9.82, which equates to around $18,400 per worker in 2018 NZD. This compares to a value-added based threshold of around $23,000 in Sin et al. (2020), using similar data.[16] Around 23% of employment is in firms with quasi-rents below this threshold, compared to 11% in Sin et al. (2020).[17] For firms with rents above this threshold the estimated sharing elasticity is 0.13, implying that a 10% increase in QR is associated with a wage increase of 1.3%.[18]

The year-specific thresholds are estimated from nominal values, so are not directly comparable across years.[19] For comparison purposes, Figure 2 plots the proportion of employment and firms in the sample with quasi-rent below the estimated thresholds. The proportion of employment in zero-rents firm varies between 15% and 30% per year, with a similar proportion of firms earning less than the threshold level of quasi-rents. On average, 23% of employment in our sample is in firms with zero excess rents. Proportions were increasing heading into the GFC. This is expected given that average quasi-rents were declining. Also, the GFC was a highly volatile and uncertain period – firms are more reluctant to give pay rises when the future state of their markets and the economy in general are uncertain. The proportion of employment in zero-rent firms has been marginally higher in the post-GFC period, compared to the pre-GFC period (21% for 2002-2007 vs. 24% for 2011-2018).

Figure A 1 in Appendix A shows the variation in the proportion of employment in zero-rent firms by industry, averaged across years. Around two-thirds of workers in the hospitality and administrative and support services work in zero-rent firms, while around one-third of workers in agriculture and retail are in zero-rent firms. The mining, finance, information media, and professional services industries all have less than 10% of workers in zero-rent firms.[20]

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[15] Table A 1 presents the estimated sharing parameters from equation 7, where the firm wage premium is used as the LHS variable. These are increasing over time, a result consistent with Allan & Maré (2021). In that paper, we found that while the relationship between wages and quasi-rents per worker is stable over time, the relative importance of worker sorting in explaining the relationship had declined, while the importance of rent sharing had increased. The importance of rent sharing was measured as the relationship between quasi-rents and firm wage premiums.
[16] Sin et al. (2020) find a lnVA based threshold of 9.95 in 2012 NZD, which equates to around $23,000 in 2018 NZD.
[17] Sin et al. (2020) look at firms with FTE>10, rather than FTE>5 as we use here.
[18] These estimates are based purely on cross-sectional variaton.
[19] Figure A 2 in Appendix A plots the estimated $\ln \tau$ over time, expressed in 2018 NZD. While there is some year-to-year volatility in the estimates, they typically fall in the range of 9.5 and 10 ($18,000-$20,000). If anything, there is a slight upward trend over time, with estimates slightly higher towards the end of the sample period.
[20] Figure 2 and Figure A 1 are based on firms with QR>0. The pattern is similar when we include firms with QR<0 as firms that earn zero rents. Between 8% and 14% of employment is in firms that earn negative quasi-rents.

There are several reasons why our estimated thresholds are significantly above zero. First, there is likely measurement error in our estimate of rents. Furthermore, in any given year a firm may earn positive rents, but there is uncertainty around whether this level of rents may be maintained into the future. This uncertainty will likely increase the threshold at which firms begin to share rents. Our measure of rents also does not account for a return to labour for any working proprietors or a 'normal' level of profit.

We place two further restrictions on our data in defining our estimation sample. We drop observations where the firm earns zero excess rents. After doing this we drop singleton observations i.e., jobs with only a single annual observation. Our estimation sample then contains 6,772,200 job-year observations on 1,107,800 individuals at 29,349 firms over 17 years.

Table 1 presents summary statistics for our estimation sample, calculated as averages across jobs. Table A 2 in Appendix A presents summary statistics for workers and firms excluded from our analysis because they earn zero excess rents or because of insufficient wage information. In Table 1, we see the average job in our sample pays around $75,000 per year and the firm earns about $78,000 in rents per worker. Approximately $59,000 of rents are excess rents per worker. The average job is in a firm with around 1300 workers and the firm is well established with an average age of 27. The proportion of jobs held by women is relatively low, at 35%. The proportion of jobs held by Māori or Pacific workers is also relatively low, at 13% and 8%, respectively.[21] The age distribution of our sample is relatively even, with between 12% and 16% of the sample in each age bracket. Compared to excluded workers, our sample is slightly older. The tenure profile of our estimation sample is skewed towards workers who have longer tenure (reflecting the exclusion of short-term jobs) and skewed towards workers with higher qualifications.

One notable feature of our sample is the relatively low percentage of women. There are two reasons for this, both related to sample selection. First, we exclude the non-market sectors: education, healthcare, and public administration. These sectors are large employers and significant employers of women. The second reason is to do with the way we calculate our wage estimates. As women work fewer hours than men on average, they are disproportionately affected by the obvious part-time exclusion.

3.2 General patterns

3.2.1 Which types of workers work in high-rent firms?

We begin by looking at which workers tend to work in high-rent firms and therefore potentially have a higher wage premium if rents are shared. Table 2 shows the average wages and quasi-rents by worker/job characteristics for our estimation sample.

[21] Māori and Pacific Peoples make up a larger proportion of workers in firms with zero excess rents and of those excluded based on insufficient wage information (see Table A 2).

The first block of Table 2 compares wage rates and rents for men and women. There is a substantial difference in the average annual wage rate between men and women in our sample. FTE earnings for men are $82,000 per year on average, compared to $61,000 for women. The unadjusted gender wage gap, based on the log wage, is 27%, similar to that found in Sin et al. (2020) using similar data.[22] While women on average earn less than men, they tend to work in firms that earn higher rents. A key reason for this somewhat surprising finding relates to our inability to distinguish high paid part-time workers from low paid full-time.

Women work fewer hours than men, on average, which means that they will have lower monthly earnings even in the absence of a gender gap in hourly wages. Our measure of part-time work is based on whether an individual earns less they would working 40 hours per week at the minimum wage. Someone working 20 hours per week for $40 per hour will look the same in our data as someone working 40 hours per week at $20 per hour, the current minimum wage. Our inability to distinguish high-paid part-time and low-paid full-time means that our estimates of the reservation wage for women (who are more likely to work part time) are artificially low. Firms with a female-dominated workforce will then have a lower estimated reservation wage bill and higher estimated rents compared to an identical male-dominated firm.

Differences in wages and rents across different groups follow expected patterns. The average wage increases with age, peaking at $81,000 in the late 40s (45-50) before steadily declining. Average rents tend to peak for slightly younger workers (30-35) and remain steady until they begin to decline for workers past the age of 45. Workers of European ethnicity have higher wages and work in firms that earn higher rents than both Māori and Pacific workers. Asian workers tend to work in higher-rent firms than Europeans but earn lower wages. Wages are increasing in the level of qualification and workers with a university education tend to work in firms that earn higher rents.[23]

3.2.2 Which types of firms earn high rents?

We next examine the firm characteristics associated with higher rents per worker. We begin this examination by looking at industry variation in average excess rents per worker. This is plotted in Figure 3, along with average wages by industry. Most industries earn between $50,000 and $100,000 in excess rents per worker. Rents are highest in the finance and insurance industry at nearly $200,000 per worker, more than double the average level shown in Table 1. Telecommunications and internet, auxiliary finance, and mining also have average rents per worker above $100,000. At the lower

[22] The official measure of the gender wage gap, based on hourly wages rather than annual earnings, in New Zealand over our sample period fluctuates between 9% and 12%, see https://www.stats.govt.nz/information-releases/labour-market-statistics-income-june-2021-quarter. The difference between our estimate and the official measure is due to differences in average hours worked between men and women.

[23] Workers with a high-school qualification are in firms that tend to earn higher rents than those with a post-school (but pre-degree) qualification. This likely reflects a difference in average age between the two groups, with workers whose highest qualification is a high-school qualification being older. Older workers had more opportunity to sort into high-rent firms.

end are grocery retailing, hospitality, and other retailing, with excess rents per worker between $9,000 and $22,000.[24] Wages are generally higher in higher rent industries, although there is much more variation in excess rents across industries than in wages.

We next explore what other firm-level factors are associated with higher (lower) rents after controlling for average differences by industry.[25] After controlling for industry differences, smaller and more capital-intensive firms tend to have larger rents, while firms with larger materials costs tend to have lower rents per worker. Firms with better quality workers, as measured by the worker fixed effect and covariate index from equation 3, have higher rents per worker despite having larger reservation wage bills. Firms with a higher pay premium have higher rents per worker. This is expected if the firm pay premium reflects the tendency to share rents.

[24] New Zealand readers may be surprised to see grocery retailing at the bottom of the rent distribution, particularly given the recent Commerce Commission market study in the grocery sector (Commerce Commission New Zealand, 2021). While it is difficult to be sure of the reason for this given the anonymised nature of the data, our hypothesis is that the firms that people think of as grocery retailers are allocated to the wholesale trade national accounts industry. The structure of the industry is that many individual supermarkets are independently owned and operated under a franchise model, with the large firms acting as the franchisor and supplier (i.e., wholesalers).

[25] We do this by regressing rents per worker on a set of firm-level characteristics: firm size, firm age, capital-labour ratio, share of intermediates in gross output, average worker quality, and the firm fixed effect from equation 3. The regression is estimated for all firms with at least 5 employees with usable wage information and observations are weighted by firm size, so the estimates reflect the experience of the average worker.

4 Estimation approach

Our aim is to estimate the extent to which rents are shared with different types of workers. We draw on the empirical framework of Card et al. (2016). Their approach was also used by Sin et al. (2020) in their exploration of the gender wage gap in New Zealand.

The starting point is an equation relating log wages to a measure of firm surplus, allowing for different groups of workers (G) to receive different shares of the surplus (rent):

$$\ln w_{ijt} = \alpha_i + X_{ijt}\beta^{G} + \gamma^{G} S_{jt} + \epsilon_{ijt} \tag{8}$$

where α_i is an individual-specific earnings premium that captures unobserved individual heterogeneity, X_{ijt} are time-varying worker and firm characteristics with (potentially) group-specific returns β^G, S_{jt} is a measure of the surplus available to be shared with workers in the form of a higher wage with group-specific sharing parameters γ^G. Card et al. (2016) further assume that S_{jt} can be expressed as:

$$S_{jt} = \bar{S}_j + \phi_{jt} + m_{ij} \tag{9}$$

where $\bar{S}_j$ captures time-invariant factors that affect the level of surplus available within a firm (e.g., market power), ϕ_{jt} captures the time variation in average surplus, and m_{ij} is the time-invariant job-specific match surplus. Substituting equation 9 into 8 and collecting terms, Card et al. (2016) then specify a variant of the standard two-way fixed effect model with group-specific firm wage premiums:

$$\ln w_{ijt} = \alpha_i + X_{ijt}\beta^{G} + \psi_j^G + e_{ijt} \tag{10}$$

where ψ_j^G are group-specific firm wage premiums and e_{ijt} is a composite error term. Wages are linked to surplus through the firm earnings premiums and the composite error term:

$$\psi_j^G = \gamma^G \bar{S}_j \tag{11}$$

$$e_{ijt} = \gamma^G\left(\phi_{jt} + m_{ij}\right) + \varepsilon_{ijt}$$

where ε_{ijt} is an idiosyncratic error term.

Card et al. (2016) estimate equation 10 separately for men and women and obtain estimates of ψ_j^G. They then relate these gender-specific firm wage premiums to time-invariant measures of surplus to obtain estimates of γ^G using the first component of equation 11.

Our interest is in documenting the extent to which firms share rents with different types of workers along several dimensions (gender, education, age etc.). Estimating group-specific firm wage premiums along multiple dimensions would be computationally intensive and time consuming. Instead, we utilise job fixed-effect models to identify the parameters of interest, γ^G.

Substituting the expressions for ψ_j^G and e_{ijt} into equation 10 yields:

$$\ln w_{ijt} = \alpha_i + X_{ijt}\beta^G + \gamma^G\left(\bar{S}_j + \phi_{jt} + m_{ij}\right) + \varepsilon_{ijt} \tag{12}$$

Including job fixed effects in equation 12 yields:

$$\ln w_{ijt} = D_{ij} + X_{ijt}\beta^G + \gamma^G\phi_{jt} + \varepsilon_{ijt} \tag{13}$$

Where $D_{ij} = \alpha_i + \gamma^G(\bar{S}_j + m_{ij})$ i.e., D_{ij} absorbs all the terms that do not vary within a job. The parameter of interest is therefore identified by the within-firm (job) time variation in measured surplus. Our empirical implementation of equation 13 substitutes our measure of $\ln EQR$ for S_{jt}. To identify group-specific sharing parameters, we interact our measure of excess rents with dummy variables for group characteristics (e.g. men and women, level of highest qualification, ethnicity, tenure, firm size). Our estimating equation is:

$$\ln w_{ijt} = D_{ij} + X_{ijt}\beta^G + \gamma^G \ln EQR_{jt} + \varepsilon_{ijt} \tag{14}$$

Where EQR is our measure of excess rents and γ^G is a vector of group-specific coefficients. Groups are defined by gender, ethnicity, highest qualification, age, and tenure. We also estimate a version where groups are defined on firm characteristics, namely firm size and firm age. To explore heterogeneity in rent sharing by industry, we estimate equation 14 separately for 39 industries.[26]

Our approach differs from the approach of Card et al. (2016) who use cross-firm variation in mean surplus to identify γ^G. Card et al. (2016) regress the estimated group-specific firm-wage premiums ψ_j^G on a time-invariant measure of surplus.[27] This approach controls for worker heterogeneity by focussing on the firm-specific components of wages. Our approach is equivalent in that we directly control for time-varying worker characteristics in the same way as the two-way fixed effect model. Furthermore, our job fixed-effects control for unobserved worker heterogeneity. They also control for permanent differences in wages and rents across firms, meaning we are identifying the parameters of interest from changes in excess rents and changes in the firm*group component of wages.

Identification

Equation 14 presents a number of identification challenges, which we discuss in Allan & Maré (2021). These include dealing with heterogeneity, measurement error and transitory shocks, and endogeneity concerns.

We use job fixed effects to control for permanent differences across workers and firms. We further control for a range of time-varying firm and worker characteristics. We include a gender-specific quartic in age, job tenure, the log of firm age, log firm FTE, the proportion of FTE within the firm with usable wage information, and the ratio of

[26] We use the pf_ind industry classification from Fabling & Maré (2015a; 2019), where each pf_ind industry is a combination of ANZSIC level 3 industries.

[27] Card et al. (2016) use average value added per worker as their measure of surplus.

intermediates to gross output. We control for macro shocks that affect both firm performance and wages by including year dummies.

Measurement error and transitory shocks are common issues when dealing with firm financial information. We expect *underlying* performance to be related to wages, but the presence of measurement error or transitory shocks means that current performance may be weakly related to underlying performance. This will cause attenuation bias in our estimates. We follow the same approach from Allan & Maré (2021), along with the bulk of the international literature, and use instrumental variables to address this.[28] In this paper, we use value added per worker as the instrument for excess rents. Value added is clearly strongly related to excess rents as it is used in the calculation. It should resolve some of the attenuation bias caused by measurement error and/or transitory shocks, particularly for the capital cost and reservation wage bill components of rents if any measurement error in value added is unrelated to measurement error in capital and the reservation wage bill. However, it will not purge quasi-rents of any measurement error in the value-added component. For this reason, our estimates of the sharing parameters should be interpreted as lower bounds.

We use the components from a 2-way fixed effect model to estimate the reservation wage to calculate EQR (see equation 5). This model contains a single firm wage premium, which is an average across the premium paid to different groups of workers at the firm. Any group-specific differences in firm pay premiums will therefore be captured by the worker fixed effects and included in the calculation of our measure of the reservation wage. Our estimate of rents will be biased downwards (upwards) if a firm's workforce contains a disproportionate number of workers who typically receive a higher (lower) share of surplus. As we are using job-fixed effects and identifying γ^G from the within-job time-series variation, permanent differences in the level of surplus are controlled for, but some of the within-job variation may be due to changing workforce composition rather than an actual change in performance.[29] This introduces a negative correlation between our measure of rents and wages, resulting in a negative bias in the estimated coefficients. Instrumenting with value added per worker, which doesn't have this mechanical relationship with workforce composition, addresses this source of bias.

Our use of a single year effect for all groups in calculating the reservation wage provides a further source of complication. The year effects capture average wage

[28] The instruments from Allan & Maré (2021) are the ratio of output to input prices measured by industry-level producer price indices, the ratio of intermediates to capital, and the ratio of intermediates to capital interacted with the industry-level produce price index for inputs. However, these instruments are relatively weak, particularly when using quasi-rents as the independent variable. See Card et al. (2018) for an overview of the international literature, including a description of instruments used in different studies.

[29] It is difficult to know how important this source of bias is as firm workforce compositions tend to be relatively stable, particularly at large firms which receive more weight in our estimation. For example, Allan & Sanderson (2021) find that the qualification structures of firm workforces are relatively stable over time, even in the presence of a major technology change.

growth across the whole economy, which we include in our measure of the reservation wage. However, wage growth has been uneven across the wage distribution, with workers in the bottom 20%-30% generally experiencing more rapid wage growth, in part due to large changes in the minimum wage over the period (Maré & Hyslop 2021). This means we are understating growth in the reservation wage for some groups of workers, and understating growth in the reservation wage bill for firms that disproportionately employ these workers. Understating growth in the reservation wage bill will then lead us to overstate growth in rents per worker.[30] This artificially strengthens the correlation between growth in rents per worker and wages, leading us to overstate the relationship between rents and wages for these groups. Instrumenting quasi-rents per worker with value added per worker eliminates this source of bias by using variation in the value-added component of quasi-rents, which is unaffected by any mismeasurement of the reservation wage.

We run regressions on the full sample and separately by gender. We do this to test the robustness of our full sample results and to test whether differential rent sharing across other characteristics (e.g., education, ethnicity) are consistent for men and women.

[30] For groups with lower-than-average wage growth, we will overstate growth in reservation wages and understate growth in rents per worker.

5 Results

We begin our analysis by presenting estimates of overall rent sharing to provide a baseline against which we can compare results from our main specification, namely job-level analysis of group-level rent-sharing elasticities (γ^G). We begin with the firm-level results from our previous work (Allan & Maré 2021) to draw the link to our previous results and to show the consistency between our previous firm-level estimates and our new job-level estimates. These are shown in Table 3.

The first four columns report firm-level firm fixed effect estimates of the relationship between quasi-rents per worker and wages, starting with all firms with positive quasi-rents (column 1). Estimates are 0.015 for the OLS estimates and 0.068 for the IV estimates. Restricting the sample in column 1 to firms with at least 5 workers with usable wage information (column 2) reduces the coefficients slightly, from 0.015 to 0.013 for the OLS results and from 0.068 to 0.046 for the IV results. IV results in columns 1 and 2 use the price-based instruments from Allan and Maré (2021). Replacing these instruments with (log) value added per worker further reduces the estimated IV coefficient from 0.046 to 0.024. However, value-added is a much stronger instrument than the price-based instruments (weak IV test statistic of 1336 vs. 3 for price-based instruments) and the standard errors in column 3 are an order of magnitude smaller than those in column 2.

Column 4 replaces quasi-rents per worker with excess rents per worker (EQR) as the main independent variable and we drop firms with zero excess rents. This increases the estimated coefficients relative to column 3, from 0.013 to 0.03 for the OLS results and 0.024 to 0.034 for the IV results. These larger coefficients are expected given the piecewise-linear relationship shown in Figure 1 and is due to dropping firms with quasi-rents below the estimated thresholds where rent sharing does not occur.[31] In column 5, we use job-level observations, use excess rents as the independent variable and exclude zero-rent firms. In this specification we replace the workforce control ($\bar{a}_i + \bar{X}_{jt}\beta$ from equation 5) with gender-specific age-earnings profiles and job tenure and use job fixed effects in place of firm fixed effects. Estimates from column 5 are slightly smaller than those in column 4, 0.028 vs. 0.030 for OLS and 0.031 vs. 0.034 for IV.[32] These estimates are our baseline and imply a 10% increase in excess rents per worker is associated with a 0.3% increase in wages, which equates to approximately a $38 increase in wages for a $1,000 in excess rents per worker.[33]

[31] Running the same regression as in column 3 but dropping firms with $\ln QR < \tau$ generates results identical to those in column 4.

[32] In our job-level estimates we are better able to control for individual heterogeneity than in the firm-level estimates, so the small reduction in point-estimates is expected.

[33] A $1,000 increase in excess rents per worker is an increase of 1.7% on the mean, equivalent to 0.014 standard deviations.

5.1 How are the rents shared with different workers?

We now examine the extent of heterogeneity in rent sharing across several worker characteristics. Table 4 presents estimated group-specific rent-sharing elasticities (γ^G). Within a column, each block of the table represents a separate regression with a different set of group interactions (e.g., male and female interactions, qualification level interactions). Column 1 reports OLS results, column 2 IV results, while column 3 converts the elasticity estimates from column 2 into the dollar change in wages associated with a $1,000 increase in excess rents per worker. Column 4 reports the results from an IV regression with all group interactions included in a single regression. Each column also reports the baseline estimate. For columns 1-3, these are from Table 3 column 5. For column 4, this is the estimate for the omitted category, which is a European male, aged 25-30, with no formal qualification and less than 1 year of tenure.

The first block considers whether rents are shared differently with men and women. OLS estimates suggest a small difference, although the confidence intervals do overlap. IV estimates are larger than the corresponding OLS estimates, and the difference between men and women is also larger. The ratio of the two IV estimates is 0.86, the wage premium from rents that women receive is 86% of that for men. Our estimate is slightly higher than the ratio found by Sin et al. (2020) of 77% and slightly lower than the 90% found in Card et al. (2016) for Portugal. In dollar terms, men receive a $49 annual wage increase for a $1,000 increase in rents per worker, while women receive $26, roughly half the amount. This is, in part, due to women having lower average wages than men, as shown in Table 2.

The second block of Table 4 looks at rent sharing for workers with different levels of qualification. The OLS point-estimates are decreasing in qualification levels, although none of the differences is statistically significant. The IV estimates are markedly different from the OLS estimates for workers with no qualifications and workers with a university qualification (bachelor's or postgraduate degree). Workers with no qualification see their wages rise by just 0.09% in response to a 10% increase in excess rents, whereas workers with a university qualification see their wages increase by 0.6%-0.7%. In dollar terms, workers with no qualifications receive $13 for a $1,000 increase in rents, compared to an increase of $92 for workers with a postgraduate degree. This could reflect the relative scarcity of workers with university qualifications, particularly advanced degrees, giving them more bargaining power and ability to secure a greater share of rents.

The third block reports results for differential sharing by ethnicity.[34] OLS estimates are reasonably similar across groups, apart from Pacific workers who, surprisingly, have a

[34] The ethnicity variables are based on a scaled total response where the sum of the ethnicity variables for each individual is equal to one. For example, if a person reports only Māori ethnicity, then the Māori variable is equal to one. If a person reports being both Māori and NZ European, these two variables will have values of 0.5. The coefficients in Table 4 reflect differences between individuals that report a single ethnicity. The rent-sharing elasticities for people that report multiple ethnicities is the average across the ethnicities they report. For

significantly higher rent-sharing elasticity. The preferred IV results present a different picture, however. In these results, Asians have the highest rent-sharing elasticity, followed by workers of European descent. IV estimates for Māori and Pacific are significantly lower than that for other groups and indistinguishable from zero. European and Asian workers receive over $40 per $1000 increase in rents, while Māori and Pacific workers receive less than $10.

The fourth block looks at rent-sharing by worker age. OLS rent sharing estimates increase with age, peaking for workers aged between 45 and 50, before declining slightly for workers aged over 55. IV estimates suggest little difference in rent sharing for workers aged between 25 and 40. IV estimates of rent sharing also peak for workers aged between 45 and 50, before declining slightly. This in part reflects longer tenure for older workers and the fact they are more likely to have moved to higher job levels within a firm and rents are more likely to be shared with these workers. It may also reflect differences in experience with bargaining across age groups. Overall, however, this age-rent sharing profile is relatively flat.

The fifth and final block considers how rent sharing differs between workers with different job tenure. Both the OLS and IV estimates show rent sharing steadily increasing with job tenure. Workers new to the firm receive a very small portion of rents, with estimated elasticities for workers with less than 2 years of tenure being below 0.01. The rent-sharing elasticity for workers with 3 or more years of tenure is 0.04. This tenure profile could reflect a learning dynamic, where both the worker and firm are learning about the quality of the job-match. It could also reflect advancement of workers within an organisation. The structure of our sample means we have at least two annual observations per job so we will see the effects of workers increasing tenure and advancement in the firm. By the time the worker has been at the firm for more than three years, they could expect to earn $50 from a $1,000 increase in rents, compared to less than $10 in their first year.

In column 4, we see that the gaps in column 2 persist even when other interactions are controlled for. The rent-sharing elasticity for women is 0.005 percentage points lower than for men, on average, and this difference is statistically significant at the 5% level. The results that the rent-sharing elasticity is increasing in both qualifications and tenure is still present in column 4, with the gaps between coefficients being slightly reduced. Accounting for differences in qualifications, age, tenure etc. reduces the difference in rent-sharing elasticities for Māori and Pacific Peoples, compared to Europeans, by around 0.5-1 percentage point, but a significant gap remains after controlling for these factors. Part of the result in column 2 is explained by differences in qualification levels (and other factors) between groups but controlling for these does not eliminate the gaps.

To gain further insights into how different workers benefit from rent sharing we estimate equation 14 for males and females separately. These results are reported in Table 5. Columns 1 and 2 report OLS estimates for males (column 1) and females

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example, a person that reports both Māori and NZ European, their estimate will be (0.5*coefficient for Māori + 0.5*coefficient for NZ European).

(column 2), while columns 3 and 4 report IV the corresponding IV results. One overall pattern is clear – the estimates for males are almost universally larger than those in the female regressions.

Some particularly large differences are apparent when looking at rent sharing with highly qualified men and women, and also in the gender-age profile of rent sharing. Men with a bachelor's degree receive a 0.7% increase in wages in response to a 10% increase in excess rents, compared to 0.4% for similarly qualified women. For those with postgraduate qualifications, men receive a 0.8% wage boost in response to a 10% increase in excess rents compared to a 0.5% increase for women. This result is likely partially driven by women (including highly qualified women) being more likely to work part time (and therefore appearing to be on lower wages), but not being captured by our obvious part-time criteria.

The age-rent sharing profile for men is relatively flat, with some indication it follows an inverted-U shape. Rent sharing increases throughout the 20s and 30s, peaking in the late 40s, before declining past the age of 50. The shape of the age profile for women is markedly different. There is relatively little difference between the rent sharing estimates for men and women aged between 25 and 35. A significant gap in rent sharing estimates opens in the late 30s. The gap narrows somewhat in the early 40s but is still significant. From the late 40s onwards, rent sharing estimates for women are similar to those for men. Differences in hours worked between men and women, and mothers and fathers in particular, is likely a key driver of this result. New Zealand research by Sin et al. (2018) shows that women work fewer hours and have lower hourly wages after having their first child, whereas there is no change in these measures for new fathers.

Other interesting gender differences are apparent, although these are smaller and imprecisely estimated. For most ethnic groups, there is a gender gap in rent-sharing elasticities of approximately one percentage point. For Māori, there is no difference between the estimates for men and women. Both coefficients are statistically insignificant. For Pacific, the difference in point estimates of 1.4 percentage points is large compared to the other groups, although imprecisely estiamted. Pacific men receive a positive, but smaller, share of rents than male workers of other ethnicities. Pacific women, on the other hand, do not appear to be sharing in any rents.

5.2 Which types of firms share more rents?

We now turn to examining whether different types of firms share their rents differently. We examine heterogeneity along three dimensions: by firm size, by firm age, and by industry.

Table 6 reports rent-sharing elasticities by firm size and age. As with Table 4, each block of the Table 6 represents a separate regression and shows the coefficients on the interaction of lnEQR and dummy variables describing the firm characteristic. Column 1 reports OLS results, column 2 IV results, while columns 3 and 4 report IV results for separate male and female regressions.

The firm-size profile of rent-sharing elasticities is relatively flat, both overall and in the separate male/female regressions. Generally mid-sized firms (between 20 and 100 FTE) have higher estimates than small (<20) and large (100+) firms. The firm-age profile is similarly flat. There is a drop in the estimated elasticity for firms aged 2-5 years. However, none of the differences in rent-sharing elasticities along firm size and age dimensions is large and are likely statistically insignificant.

Figure 4 plots rent sharing estimates from industry-specific IV regressions.[35] Industries are ordered by their average excess rents per worker as in Figure 3. Also shown is the 95% confidence interval for the overall rent-sharing elasticity from column 5 of Table 3 (dashed lines Figure 4).

Three broad groups of industries are evident in the figure. The first are the low-rent industries, from grocery retailing to administrative and support services. Despite these industries having relatively low rents, the rent-sharing elasticities are comparable to the overall elasticity shown in Table 3. The point estimates are generally increasing in the level of rents (moving left to right in the figure), although these differences are generally not statistically significant.

Part of the reason for the comparatively high rent-sharing elasticities for these low-rent industries is how we have selected our sample. Many of these industries have a high proportion of employment in zero-rent firms, meaning they are excluded from our analysis. These industries also have a relatively high proportion of part-time and/or short-term workers, who are excluded from the analysis because of insufficient wage information. Remaining full-time workers are more likely to be in management or supervisory type roles. We are therefore estimating the rent-sharing elasticity for workers who may be more likely to receive a share of rents at the relatively small group of firms that have sufficient rents to share.

The second group of industries are roughly in the middle of the excess rent distribution (heavy & civil construction to wholesale trade). There is a noticeable step down in the estimated rent-sharing elasticities compared to the previous group of industries, with many industries in this group having lower estimates compared to administrative services, printing, and services to the primary sector. Estimates in this group are similar and don't increase with the level of excess rents (moving left to right). One exception is transport equipment manufacturing. This industry has one of the larger point estimates at 0.05.

The final group of firms (other transport to finance and insurance) are distinguished by the degree of heterogeneity in rent sharing estimates across these industries. These industries have the highest levels of average excess rents. Industries such as food and beverage manufacturing and utilities have very small estimates of the rent-sharing elasticity that are statistically indistinguishable from zero. Others, such as professional and technical services and auxiliary finance (mostly brokering services) have some of the highest rent-sharing elasticities. Some of these industries are dominated by large, profitable companies with a diverse workforce, while others are dominated by smaller

[35] OLS estimates are shown as hollow circles.

firms. Some workers in the large, profitable firms may receive a disproportionate share of rents, but most workers may receive little benefit from improving firm performance, resulting in a lower rent-sharing elasticity.

The somewhat negative relationship between rent sharing and the level of rents across industries reflects the fact that the differences in rents between industries are significantly larger than differences in wages. Rents per worker differ by a factor of 20 when comparing the highest rent industry to the lowest. The same comparison for wages is a factor of nearly 2.5, a vast disparity. A 10% increase in rents in the most profitable industries is a far larger absolute dollar increase than a 10% increase in the least profitable. While it appears that workers in low-rent industries receive a larger share of rents, this is a larger share of a smaller pie.

To get a better sense of sharing across industries, we convert the rent-sharing elasticities to look at sharing by cents in the dollar. This is plotted in Figure 5. As with the elasticity estimates, sharing in per dollar terms is generally higher in industries with lower rents, and estimated sharing is decreasing in the level of rents. Notable exceptions are the professional, scientific, and technical services, and auxiliary finance industries. Firms at the bottom end of the rent distribution share more of their rents, between 5 and 12 cents per dollar, but simply have less rent to share. An increase of excess rents per worker of $1,000 represents a 11% increase in rents for the average firm in the grocery retailing industry, and a 5% increase for the average hospitality firm. At the other end of the distribution, $1,000 represents a 0.5% increase in rents per worker in the finance and insurance industry, a 0.7% increase in the mining industry, or a 1% increase in the telecommunications and internet industry. While these higher-rent industries share less per dollar (between 1 and 4 cents per dollar), there are many more dollars to share.

Figure 6 converts these per-dollar rent sharing estimates into total amount of rents received by the average worker per year by multiplying the estimates from Figure 5 by average excess rents per worker in each industry. While sharing per dollar across industries tends to be lower in higher rent industries, workers in different industries tend to receive a similar amount, on average. For most industries, workers receive between $1,500 and $2,000 a year from rents. Workers in the auxiliary finance industry receive nearly $7,000 per year and those in the professional, technical, and scientific services $4,000 per year. At the other end of the spectrum are workers in the grocery retailing, food and beverage manufacturing, and utilities sectors, who receive less than $1,000 per year in rents.

Another reason that rent sharing may differ across industries is that some industries may face demand conditions that are highly changeable, leading to higher uncertainty about future firm performance. This uncertainty about demand conditions (and therefore future rents) may induce some insurance-type behaviour, where firms bank most of the excess rents in good years to keep wages steady in bad years. If this were the case, then we would expect to see a negative relationship between volatility in rents (a proxy for uncertainty) and rent-sharing elasticities across industries.

Figure 7 plots the estimated rent-sharing elasticities (from the IV model) against the industry-average within-firm standard deviation in (log) excess rents per worker. This provides us a measure of how much uncertainty in rents per worker the firm of the average worker faces.

We see a slight negative relationship between the estimated rent-sharing elasticities and average within-firm standard deviations. This suggests that some of the differences across industries may reflect greater insurance behaviour. If certain groups of workers are overrepresented in certain industries, then part of the differences between groups may also reflect insurance behaviour. However, the correlation between the estimated rent-sharing elasticities and within-firm standard deviations is relatively weak (correlation coefficient -0.11, regression coefficient of -0.024, t-statistic of -0.74), suggesting other factors are also at play. However, two industries could be considered outliers with either very low within-firm standard deviations (grocery retailing) or negative rent-sharing elasticity (other transport). Excluding these industries strengthens the correlation between the rent sharing elasticities and within-firm standard deviations (correlation coefficient -0.38, regression coefficient -0.07, t-statistic -2.37).

It is difficult to draw strong conclusions regarding differences in rent sharing between firms with different characteristics (age, size, industry). Differences in estimated coefficients are relatively small and standard errors sufficiently large, making it difficult to say that workers in different types of firms are more successful at obtaining a share of rents.

6 Conclusion and discussion

In this paper we explore heterogeneity in rent sharing by a range of worker and firm characteristics. This provides insights into the role of firms in explaining a range of wage inequalities between different groups in New Zealand.

We use a measure of excess quasi-rents using the method of Card et al. (2016). We find that the relationship between quasi-rents and the firm component of wages is flat for firms that earn less than $18,000-$20,000 rents per worker in 2018 NZD. Around 23% of workers are in firms that earn no excess rents and therefore do not benefit from rent sharing. These workers are concentrated in the hospitality, administrative services, agriculture, and retail industries. These workers are also more likely to be women and more likely to identify as Māori or as Pacific Peoples. They are also younger and have lower rates of higher education. Institutional factors, such as changes to statutory minimum wages, are significant drivers of wage growth at these firms.

We find an overall rent-sharing elasticity of 0.03, which is at the lower end of comparable international studies (see Card et al. 2018). However, in the Card et al. (2016) framework, on which we base our empirical strategy, there is a group-specific sharing parameter that applies to total surplus. This may not necessarily be the case in practice, particularly if firms provide insurance against temporary swings in firm performance. Studies examining insurance-type behaviour by firms find that firms provide greater insurance against temporary swings in performance than permanent change in performance. The true sharing parameter that applies to the within-firm variation in surplus may be lower than that applying to the cross-firm variation if firms insure their workers against temporary swings in performance and temporary shocks are a significant component of within-firm variance.

For workers in firms earning positive excess rents, the results suggest that some groups of workers benefit more from improvements in firm performance than others. Men earn greater rents than women, consistent with Sin et al. (2020). Workers with higher qualifications and workers with longer tenure experience greater proportionate wage increases in response to an increase in firm performance. Conversely, Māori and Pacific workers, in particular Pacific women, benefit less from increases in rents per worker. Their wages are still likely to be increasing (e.g., increases in the minimum wage), but this has less to do with the performance of the firm at which they work. Overall, our results are consistent with aggregate statistics on wage gaps between different groups and provide further evidence that within-firm gaps contribute to the aggregate wage gaps.

It is difficult to draw strong conclusions on the relative success of workers obtaining a share of rents in different industries, given the relative imprecision of the industry-specific estimates. Industries with lower average rents per worker tend to have higher rent-sharing elasticities and a higher rate of sharing in cents per dollar. This pattern is partly to do with selection - a large fraction of workers in these industries are in firms that do not share any rents (rents per worker less than the threshold level) and a large proportion of workers are part time. The relatively high rent-sharing elasticities therefore reflect the sharing at a small proportion of firms that earn sufficient rents and the small proportion of workers that are possibly more likely to receive a share. While the elasticities and sharing per dollar are larger in these industries, the average amount shared with workers is similar, consistent with these workers getting a larger share of a smaller pie.

There are several reasons why some groups benefit more (less) from increases in rents per worker. An obvious consideration is differences in the relative supply and demand for workers with different skills. It is also well established that different groups (e.g., men and women) have different attitudes towards wage bargaining (Niederle & Vesterlund 2011). This could be driving the gender difference in rent sharing and may also partially explain the differences in rent sharing by ethnicity as well (The Auckland Co-Design Lab 2016; Equal Employment Opportunities Trust 2011). Another potential explanation is differences in attitudes to hierarchy between groups. Some groups may be less likely to raise issues or ask for higher wages from their employers. Differences across groups could also reflect differences in relative positions within a firm. For instance, longer tenure workers are more likely to have moved up the job ladder within a firm and therefore better able to obtain rents. Similarly, workers with higher qualifications may be more likely to be in management, supervisory, or specialist roles. Finally, it could reflect differences in union or collective bargaining coverage across groups or industries.

In addition, firms may possess greater monopsony power over some workers. Different workers may have different preferences for hours of work, work location and commutes, preferences for work versus leisure, or other firm attributes that affect the relative attractiveness of a particular firm as a workplace. These differing preferences mean that firms may not need to raise wages as much to dissuade some workers from leaving or may have to offer large wage increases to entice workers from other firms. Some groups may have greater difficulty finding a new job, which could reflect differences in their geographic distribution. Different groups may also have different attitudes to discussing pay and may therefore have less information on wages offered at other firms, or even to other workers at the same firm. All these factors could contribute to differences in the amount of monopsony power that firms possess over different workers.

We find suggestive evidence that some of the differences across groups may reflect insurance type behaviour. We find a negative correlation between industry-specific rent-sharing elasticities and within-firm standard deviations of excess rents, a proxy for the degree of uncertainty about future rents. Where there is greater uncertainty, we would expect firms to retain most of the rents from a good year to help them navigate tougher economic times, resulting in a lower rent-sharing elasticity.

This study, along with Sin et al. (2020), are the first New Zealand studies to examine the role of rent sharing in explaining differences in wages between groups. We expand on their results and show that differences in rent sharing are important for explaining ethnic wage gaps as well. While we have speculated on the reasons for such differences and provided suggestive evidence consistent with insurance being a partial explanation, further work is needed to dig more deeply into the reasons behind these differences in rent sharing with the aim to informing policy design to reduce these inequalities in New Zealand society.

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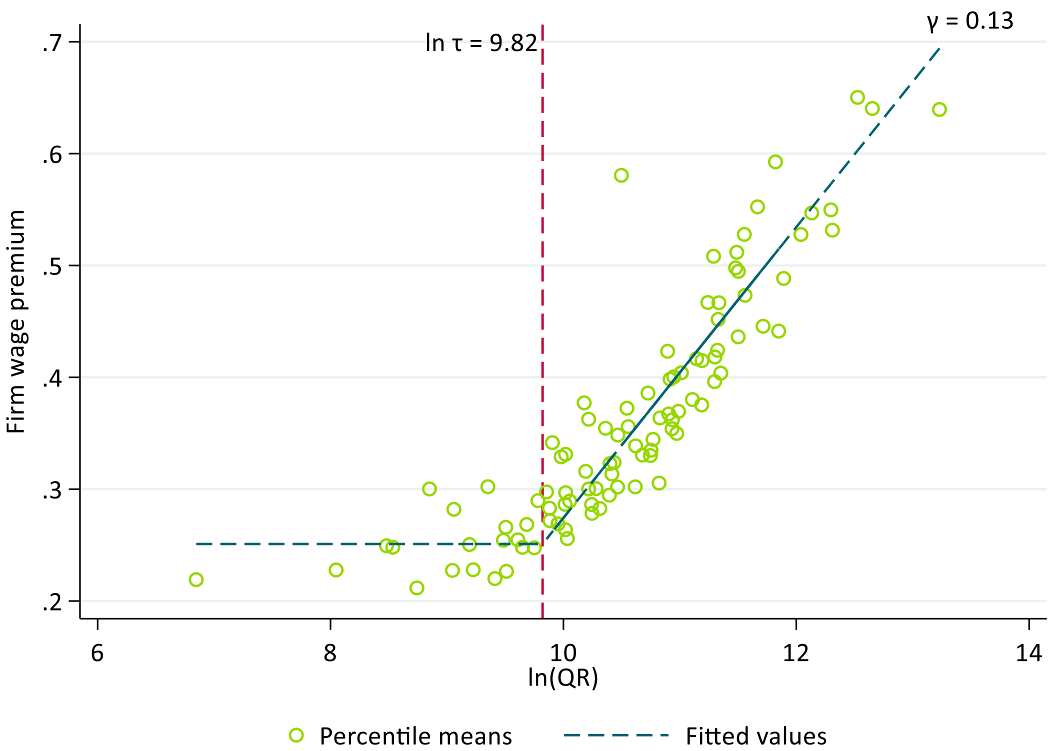
Figures

Figure 1: The piecewise linear relationship between quasi-rents per worker and firm wage premiums, 2018



Notes: Each point represents the average firm wage premium and quasi-rents per worker in 100 percentile bins in the quasi-rents per worker distribution. The blue dashed lines represent fitted values, estimated by non-linear least squares (NLS). The red vertical dashed line is the NLS estimate of τ.

Figure 2: Percent of firms and employment that earn zero excess quasi-rents

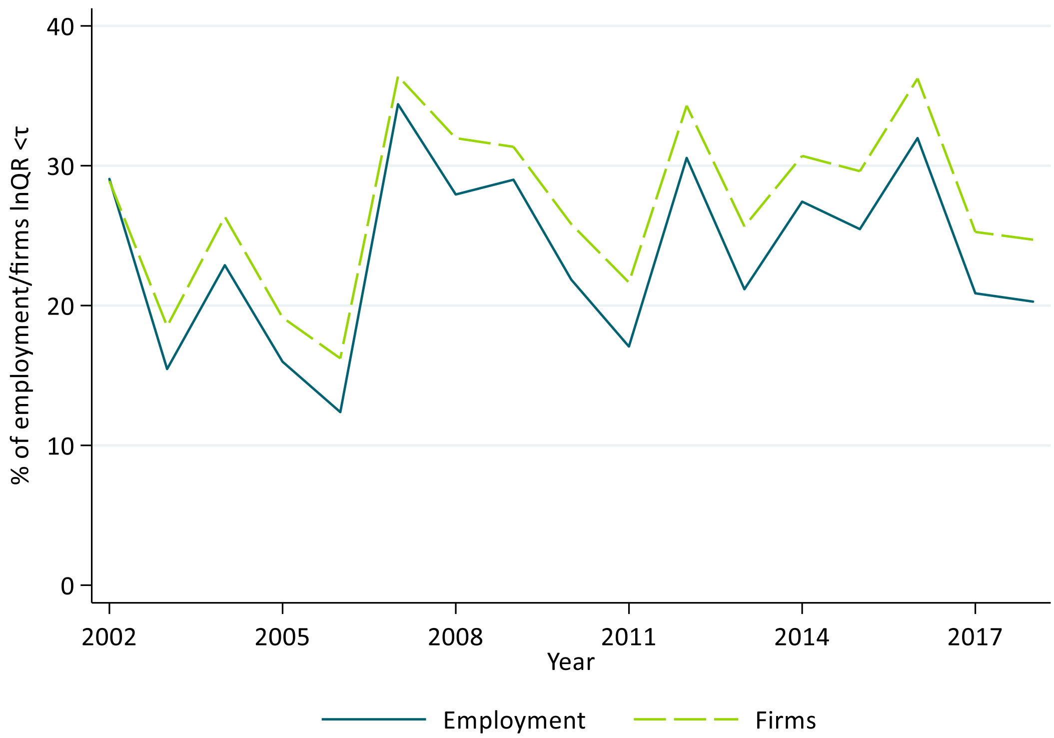


Figure 3: Excess rents per worker and wages by industry

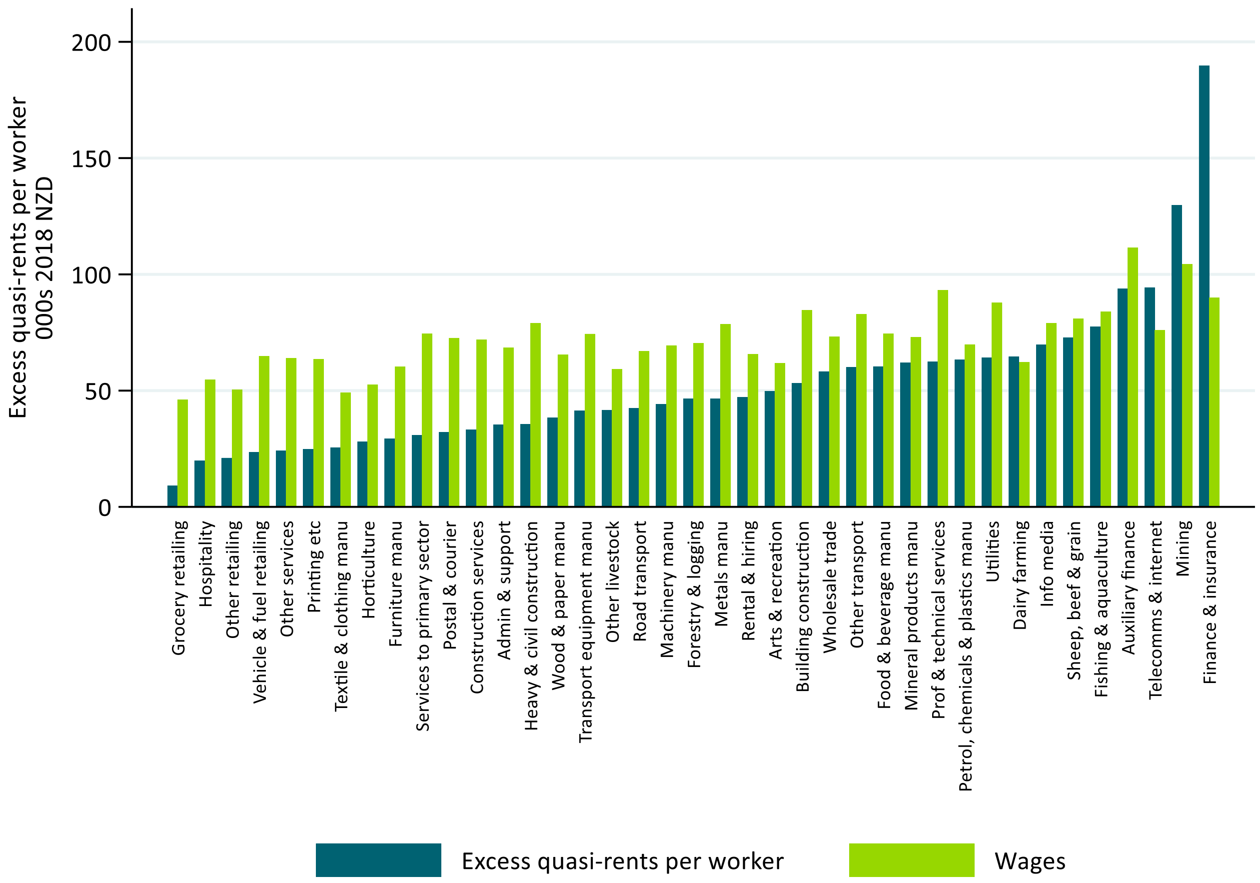
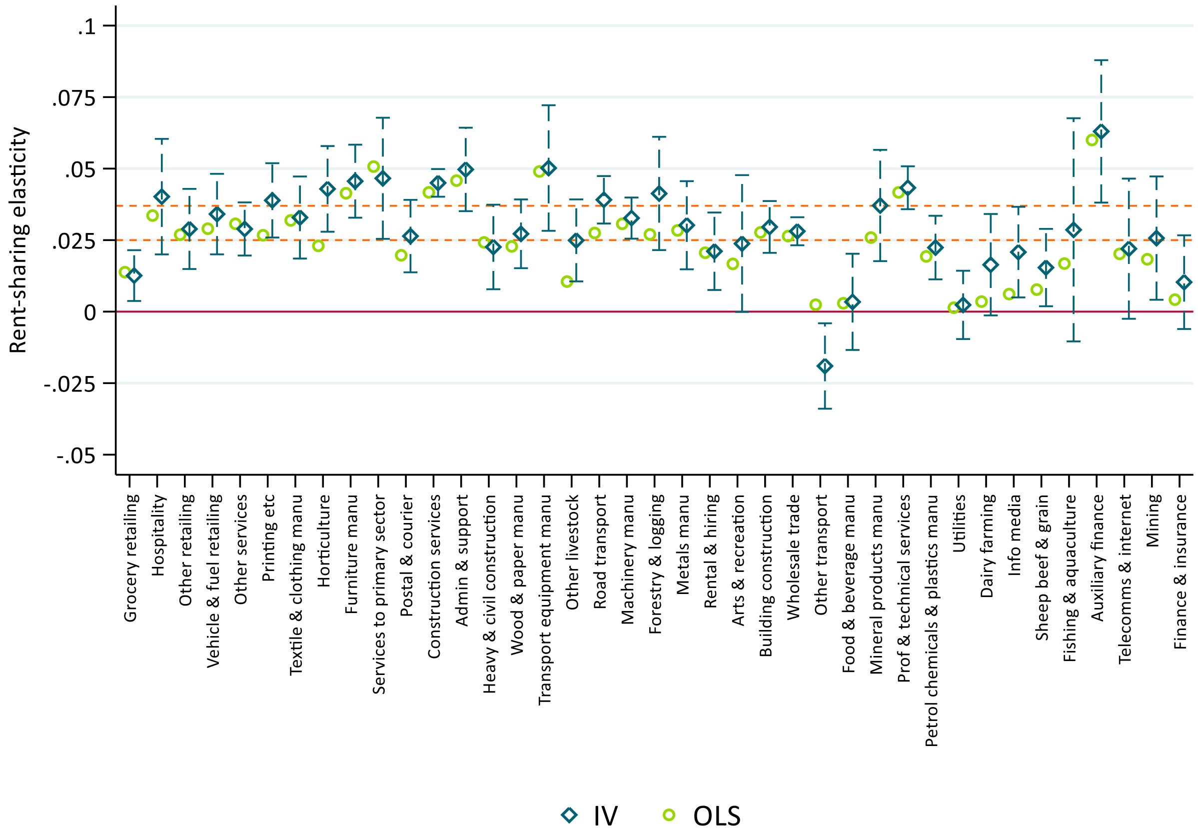


Figure 4: Rent-sharing elasticities by industry



Notes: Diamonds represent industry-specific IV estimates of equation 14. Circles represent the corresponding OLS estimates. Dashed lines are the upper and lower bounds of the 95% confidence interval for the IV estimate of overall rent sharing (Table 3 column 5).

Figure 5: Rent sharing per dollar by industry

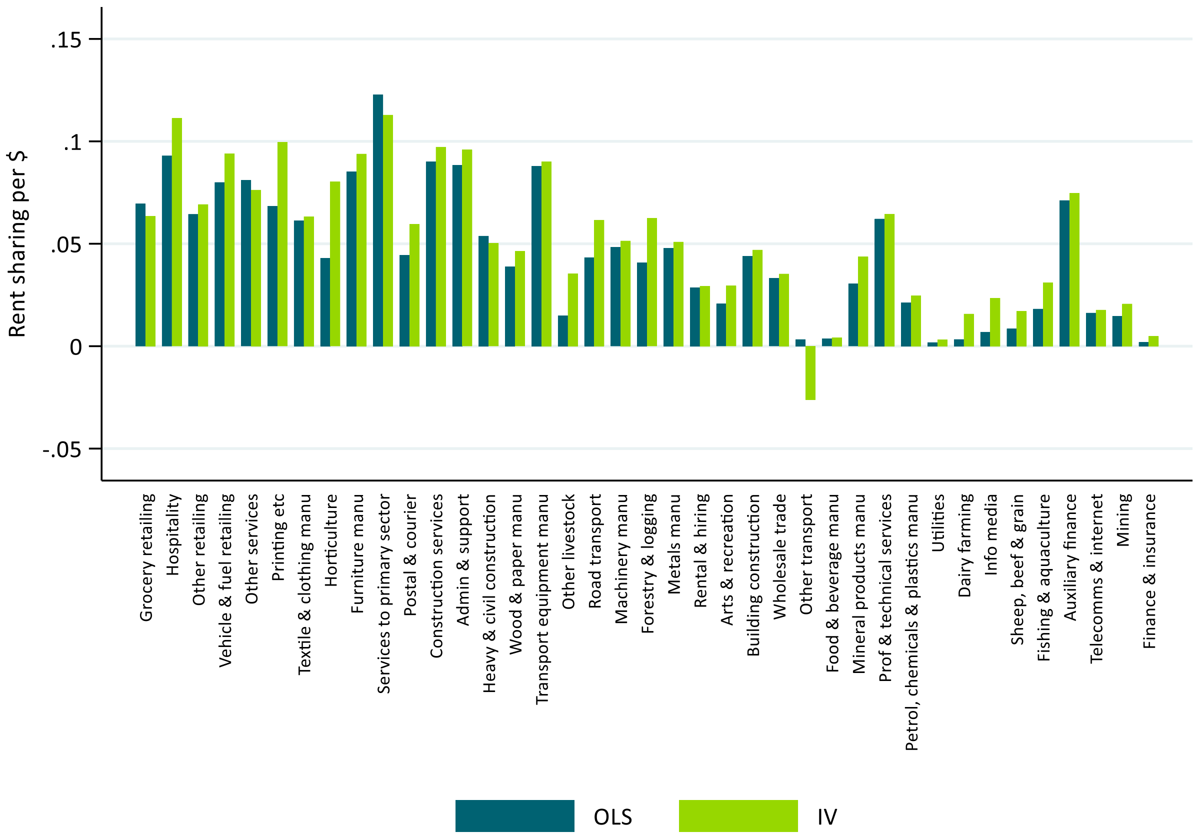


Figure 6: Average amount of rents received per worker per year by industry

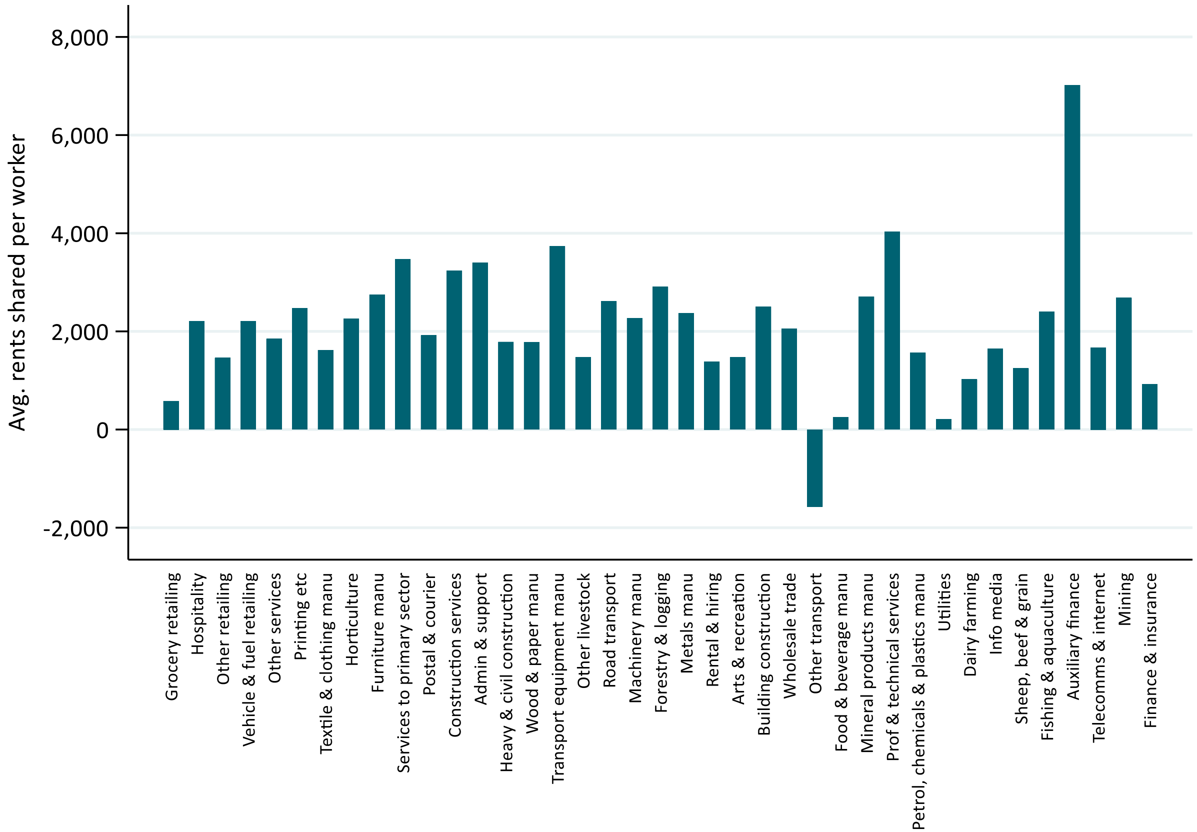
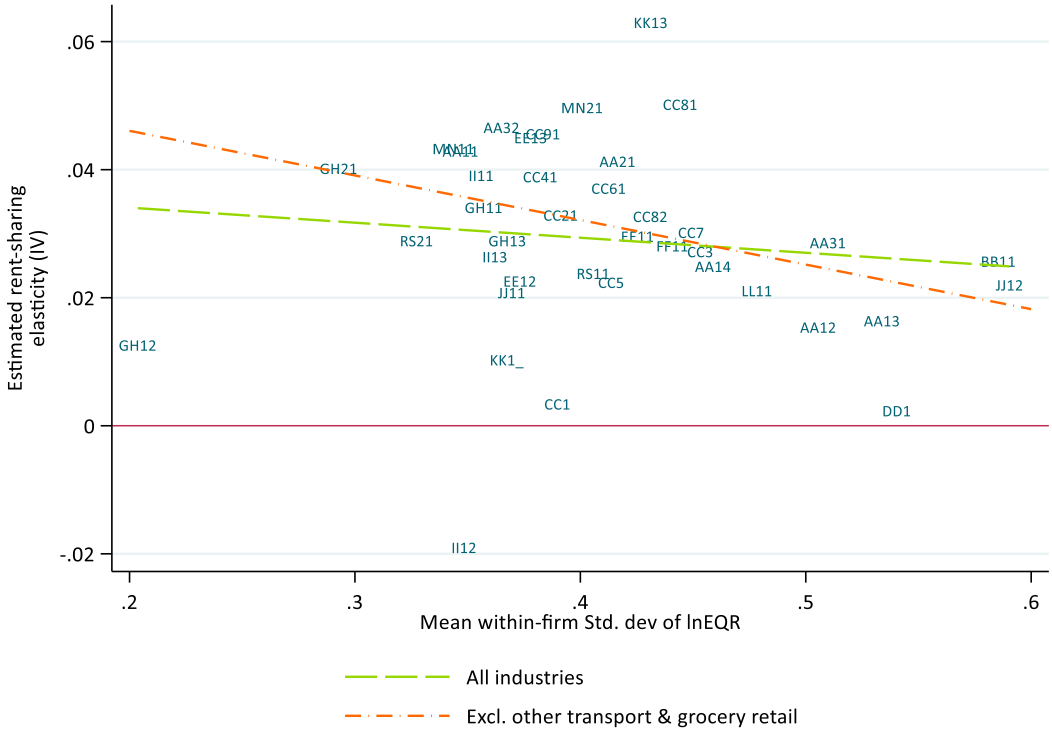


Figure 7: Rent sharing and firm-level volatility in rents per worker



Notes: Industry codes mark the estimated rent-sharing elasticity and mean within-firm standard deviation for the industry. Industry codes are: AA11 Horticulture; AA12 Sheep, beef & grain farming; AA13 Dairy farming; AA14 Other livestock; AA21 Forestry & logging; AA31 Fishing & aquaculture; AA32 Services to primary sector; BB11 Mining; CC1 Food & beverage manufacturing; CC21 Textile & clothing manufacturing; CC3 Wood & paper manufacturing; CC41 Printing etc; CC5 Petrol, chemicals & plastics manufacturing; CC61 Mineral products manufacturing; CC7 Metals manufacturing; CC81 Transport equipment manufacturing; CC82 Machinery manufacturing; CC91 Furniture manufacturing; DD1 Utilities; EE11 Building construction; EE12 Heavy & civil construction; EE13 Construction services; FF11 Wholesale trade; GH11 Vehicle & fuel retailing; GH12 Grocery retailing; GH13 Other retailing; GH21 Hospitality; II11 Road transport; II12 Other transport; II13Postal & courier; JJ11 Info media; JJ12 Telecommunications & internet; KK13 Auxiliary finance; KK1_ Finance & insurance; MN11 Prof & technical services; MN21 Admin & support services; RS11 Arts & recreation; RS21 Other services

Tables

Table 1: Summary statistics for estimation sample – jobs in firms with positive excess rents

Wages and firm chars		Demographics	
Real annual wage	\$74,700	Female	.352
	(\$49,400)		(.478)
Log real annual wage	11.097	European	.758
	(.462)		(.428)
Real quasi-rents pw	\$77,800	Māori	.129
	(\$72,700)		(.335)
Log real quasi-rents pw	10.969	Pacific	.079
	(.724)		(.27)
Real excess rents ($QR - e^{\ln \tau}$)	\$59,200	Asian	.108
	(\$72,500)		(.31)
Excess log rents ($\ln QR - \ln \tau$)	1.158	MELAA	.013
	(.732)		(.112)
Employment (FTE)	1348	Missing ethnicity information	.004
	(2455)		(.064)
Log employment (FTE)	5.397	Age 25-30	.139
	(2.123)		(.346)
Firm age	27	Age 30-35	.15
	(23)		(.357)
		Age 35-40	.15
			(.357)
		Age 40-45	.15
			(.357)
		Age 45-50	.138
			(.345)
		Age 50-55	.115
			(.319)
		Age 55+	.159
			(.366)
		No qualifications	.143
			(.35)
		High-school qualifications	.323
			(.468)
		Post-school qualifications	.229
			(.42)
		Bachelor's	.125
			(.331)
		Postgraduate qualifications	.081
			(.273)
		Missing qualifications	.099
			(.298)

Table 1 cont.

	Tenure < 1 year[36]	.047
		(.213)
	Tenure 1-2 years	.15
		(.358)
	Tenure 2-3 years	.188
		(.391)
	Tenure 3+ years	.614
		(.487)
N		6,772,200
N individuals		1,107,800
N Firms		29,349
Avg. obs per job		4.42

Notes: The number of observations and number of individuals have been graduated random rounded for confidentiality purposes. The number of firms has been randomly rounded to base 3 for confidentiality purposes. Standard deviations are in parentheses. The ethnicity variables are based on total response so will sum to greater than 1.

[36] As we are using administrative tax data, our only information on job tenure is the number of months we observe a person working at a particular firm. The tax data begins in 1999 and our sample begins in 2002, meaning that our estimate of tenure is censored at 3 years in the first year of our sample. This is the reason we have specified our tenure variable as a set of binary indicators, with the long tenure category being 3+ years.

Table 2: Average wages and rents by worker/job characteristics

	N indiv	N firms	Wage	lnWage	QR	lnQR	VA	lnVA	K/L ratio
				Gender					
Male	701,000	28,539	$81,900	11.19	$73,500	10.94	$152,300	11.82	$29,600
			($26)	(.0002)	($32)	(.0003)	($42)	(.0002)	($20)
Female	406,700	25,200	$61,300	10.92	$85,900	11.03	$157,200	11.83	$26,600
			($23)	(.0003)	($53)	(.0005)	($63)	(.0003)	($25)
				Age					
25-30	336,900	24,819	$59,700	10.94	$79,800	10.99	$153,600	11.82	$27,400
			($25)	(.0004)	($76)	(.0008)	($95)	(.0005)	($41)
30-35	369,000	25,902	$70,200	11.07	$81,500	11.01	$157,300	11.84	$28,500
			($34)	(.0004)	($75)	(.0007)	($94)	(.0005)	($41)
35-40	358,700	25,782	$77,500	11.14	$81,100	11	$157,900	11.84	$29,000
			($46)	(.0005)	($75)	(.0007)	($95)	(.0005)	($42)
40-45	349,300	25,437	$81,000	11.16	$79,000	10.98	$156,000	11.83	$28,900
			($56)	(.0005)	($73)	(.0007)	($93)	(.0005)	($41)
45-50	320,000	24,618	$81,300	11.16	$77,000	10.96	$153,800	11.82	$28,700
			($61)	(.0005)	($75)	(.0007)	($94)	(.0005)	($41)
50-55	268,100	23,058	$79,700	11.14	$74,900	10.94	$151,700	11.81	$28,700
			($68)	(.0006)	($79)	(.0008)	($101)	(.0005)	($45)
55+	239,900	22,590	$73,900	11.08	$71,300	10.9	$147,800	11.79	$28,500
			($51)	(.0005)	($64)	(.0007)	($82)	(.0004)	($37)
				Ethnicity					
European	814,100	28,797	$78,700	11.14	$79,200	10.99	$156,200	11.83	$28,700
			($24)	(.0002)	($33)	(.0003)	($41)	(.0002)	($18)
Māori	148,300	19,683	$64,200	10.99	$68,000	10.87	$144,600	11.77	$29,300
			($34)	(.0004)	($67)	(.0007)	($90)	(.0005)	($41)
Pacific People	87,200	11,703	$57,900	10.9	$69,700	10.88	$143,000	11.75	$28,300
			($33)	(.0005)	($89)	(.0009)	($116)	(.0006)	($50)
Asian	137,300	14,268	$65,600	11	$84,400	11.01	$157,900	11.84	$27,600
			($40)	(.0005)	($94)	(.0009)	($110)	(.0006)	($45)
MELAA	19,200	6,819	$70,200	11.06	$77,000	10.98	$153,600	11.83	$29,200
			($144)	(.0015)	($239)	(.0024)	($318)	(.0016)	($152)
Missing	8,300	3,672	$85,100	11.14	$73,300	10.92	$147,300	11.78	$25,800
			($515)	(.0034)	($397)	(.0042)	($513)	(.0028)	($219)
				Qualifications					
None	144,500	20,037	$58,900	10.92	$61,000	10.78	$135,100	11.71	$27,900
			($25)	(.0004)	($55)	(.0007)	($76)	(.0004)	($33)
High school	331,500	27,198	$66,600	11.01	$78,100	10.96	$152,400	11.81	$28,100
			($26)	(.0003)	($50)	(.0005)	($62)	(.0003)	($26)
Post-school	241,500	25,941	$76,400	11.15	$72,200	10.92	$150,100	11.8	$29,000
			($32)	(.0003)	($53)	(.0005)	($71)	(.0004)	($35)
Bachelor’s	142,400	17,670	$93,100	11.28	$97,800	11.18	$176,400	11.95	$29,400
			($73)	(.0006)	($95)	(.0008)	($113)	(.0005)	($49)
Postgrad	91,600	12,852	$107,900	11.43	$100,300	11.23	$181,800	11.98	$30,100
			($105)	(.0007)	($116)	(.001)	($143)	(.0007)	($67)
Missing	156,300	20,289	$69,300	11.02	$70,700	10.9	$144,900	11.77	$27,200
			($64)	(.0006)	($79)	(.0008)	($104)	(.0006)	($46)

Table 2 cont.

				Tenure					
< 1 year	250,600	20,646	$66,900	11	$79,500	11	$155,200	11.83	$27,800
			($72)	(.0008)	($130)	(.0013)	($164)	(.0008)	($69)
1-2 years	677,400	27,429	$68,100	11.02	$77,300	10.98	$152,500	11.81	$27,300
			($42)	(.0004)	($70)	(.0007)	($89)	(.0005)	($39)
2-3 years	799,600	28,185	$69,100	11.03	$74,100	10.94	$149,000	11.79	$27,300
			($38)	(.0004)	($61)	(.0006)	($78)	(.0004)	($35)
3+ years	867,400	28,701	$78,600	11.14	$79,000	10.98	$155,900	11.83	$29,300
			($26)	(.0002)	($36)	(.0004)	($46)	(.0002)	($20)

Notes: Individual and firm counts have been subjected to graduated random rounding and random rounding to base 3, respectively. Statistics by ethnicity are based on total response e.g., if a person identifies as European and Māori, they will be counted in both groups. MELAA stands for Middle Eastern, Latin American, and African. Standard errors are presented in parentheses. K/L ratio is the capital labour ratio.

Table 3: The relationship between rents and wages - moving from the firm-level to the job-level

	1	2	3	4	5
	All QR>0[37]	QR>0 and FTE>5	QR>0 and FTE>5 – VApw as instrument	EQR>0	Job level – EQR>0
			OLS		
lnQR	0.015***	0.013***	0.013***		
	(0.001)	(0.001)	(0.001)		
lnEQR				0.030***	0.028***
				(0.002)	(0.002)
R2	0.971	0.980	0.980	0.979	0.952
			IV		
lnQR	0.068***	0.046***	0.024***		
	(0.026)	(0.012)	(0.002)		
lnEQR				0.034***	0.031***
				(0.003)	(0.003)
R2	0.128	0.250	0.341	0.381	0.089
Fixed effects	Firm, Year	Firm, year	Firm, Year	Firm, Year	Job, Year
Control vars	Firm, workforce	Firm, workforce	Firm, workforce	Firm, workforce	Firm, demographics
Instruments	Prices	Prices	VA	VA	VA
Under-id	0.000	0.000	0.000	0.000	0.000
Weak IV	3.432	3.050	1336	3130	3882
Over-id	0.158	0.044	-	-	-
N	1,072,656	262,806	262,806	188,682	6,772,200
N Firms	170,754	37,254	37,254	29,472	29,349

Notes: Standard errors in parentheses are clustered at the firm level. The number of observations and number of firms have been randomly rounded for confidentiality purposes. Under-id is p-value for the Kleibergen-Paap rk LM test for under-identification. Over-id is the p-value for the Hansen J test of overidentifying restrictions. Weak IV is the Kleibergen-Paap Wald rk F statistic for the test of weak instruments. Equations are estimated using the reghdfe (OLS) and ivreghdfe (IV) commands in Stata16 (Correia, 2017; Baum, Schaffer, & Stillman, 2010).

[37] These results are based on the same specification as in Allan and Maré (2021) Table 3, column 5. The OLS estimate differs slightly from that reported in our previous work. The results reported here are based on observations with positive quasi-rents only.

Table 4: Heterogeneous rent-sharing - how rents are shared with different workers

		1 OLS	2 IV	3 IV ($ wages per $1000 of rents)	4 IV – all interactions
Baseline		0.028***	0.031***	$38	-0.028***
		(0.002)	(0.003)		[0.005]
Gender	*Male*	0.029***	0.033***	$49	-
		(0.002)	(0.003)		
	Female	0.026***	0.028***	$26	-0.005**
		(0.003)	(0.003)		[0.002]
	R^2	0.952	0.089		
	Under-id	-	179.6		
	Weak IV	-	2927		
Qualifications	*No quals*	0.030***	0.009***	$13	-
		(0.002)	(0.003)		
	High school	0.029***	0.022***	$24	0.010***
		(0.002)	(0.002)		[0.002]
	Post-school	0.028***	0.030***	$42	0.017***
		(0.002)	(0.002)		[0.004]
	Bachelor's	0.027***	0.060***	$71	0.046***
		(0.004)	(0.007)		[0.008]
	Postgraduate	0.025***	0.070***	$92	0.056***
		(0.005)	(0.010)		[0.011]
	R^2	0.952	0.086		
	Under-id	-	287.3		
	Weak IV	-	248.7		
Ethnicity	*NZ European*	0.030***	0.035***	$46	-
		(0.002)	(0.003)		
	Maori	0.026***	0.002	$3	-0.024***
		(0.003)	(0.005)		[0.005]
	Pacific Peoples	0.038***	0.007	$8	-0.024***
		(0.002)	(0.004)		[0.004]
	Asian	0.030***	0.041***	$41	-0.006
		(0.003)	(0.007)		[0.006]
	MELAA	0.025***	0.030***	$37	-0.013***
		(0.003)	(0.005)		[0.004]
	R^2	0.952	0.088		
	Under-id	-	241.7		
	Weak IV	-	171.7		

Table 4 cont.					
Age	*25-30*	0.011***	0.029***	$28	-
		(0.002)	(0.003)		
	30-35	0.022***	0.028***	$31	-0.001**
		(0.002)	(0.003)		[0.0005]
	35-40	0.026***	0.028***	$34	-0.002**
		(0.002)	(0.003)		[0.001]
	40-45	0.032***	0.032***	$43	0.002***
		(0.002)	(0.003)		[0.001]
	45-50	0.036***	0.036***	$50	0.006***
		(0.002)	(0.003)		[0.001]
	50-55	0.034***	0.034***	$48	0.004***
		(0.0025)	(0.0026)		[0.001]
	55+	0.031***	0.030***	$43	0.000
		(0.003)	(0.003)		[0.001]
	R^2	0.952	0.090		
	Under-id	-	150.6		
	Weak IV	-	109.9		
Tenure	*<1 year*	0.005	0.001	$10	-
		(0.004)	(0.004)		
	1-2 years	0.001***	0.001***	$11	0.010***
		(0.002)	(0.003)		[0.003]
	2-3 years	0.018***	0.018***	$23	0.020***
		(0.002)	(0.002)		[0.003]
	3+ years	0.034***	0.039***	$50	0.041***
		(0.002)	(0.003)		[0.004]
	R^2	0.952	0.092		0.081
	Under-id	-	90.6		154
	Weak IV	-	307.4		38
N		6,772,200	6,772,200	6,772,200	6,772,200
N indiv		1,107,800	1,107,800	1,107,800	1,107,800
N firms		29,349	29,349	29,349	29,349

Notes: See notes to Table 3. In column 4, the first entry in each block of the table is the omitted category, and the remaining estimates are relative to the omitted category.

Table 5: Heterogeneous rent-sharing - gender-specific regressions

		1	2	3	4
		OLS	OLS	IV	IV
		Males	Females	Males	Females
Qualifications	*No quals*	0.032***	0.023***	0.009***	0.010***
		(0.003)	(0.003)	(0.003)	(0.004)
	High school	0.032***	0.023***	0.022***	0.021***
		(0.002)	(0.003)	(0.002)	(0.003)
	Post-school	0.030***	0.022***	0.031***	0.024***
		(0.003)	(0.003)	(0.003)	(0.003)
	Bachelor's	0.028***	0.024***	0.071***	0.041***
		(0.005)	(0.004)	(0.009)	(0.006)
	Postgraduate	0.026***	0.022***	0.078***	0.052***
		(0.006)	(0.006)	(0.011)	(0.010)
	R^2	0.954	0.938	0.098	0.066
	Under-id	-	-	256.998	182.779
	Weak IV	-	-	316.956	93.821
Ethnicity	*NZ European*	0.030***	0.022***	0.038***	0.029***
		(0.00266)	(0.003)	(0.003)	(0.003)
	Maori	0.028***	0.018***	0.000	0.006
		(0.003)	(0.004)	(0.006)	(0.006)
	Pacific	0.042***	0.028***	0.012***	-0.002
		(0.003)	(0.005)	(0.004)	(0.007)
	Asian	0.033***	0.026***	0.045***	0.035***
		(0.004)	(0.005)	(0.007)	(0.009)
	MELAA	0.028***	0.017***	0.032***	0.028***
		(0.004)	(0.006)	(0.006)	(0.007)
	R^2	0.954	0.938	0.102	0.066
	Under-id	-	-	385.258	85.815
	Weak IV	-	-	257.523	74.052
Age	*25-30*	0.008***	0.013***	0.028***	0.029***
		(0.003)	(0.004)	(0.003)	(0.003)
	30-35	0.021***	0.020***	0.029***	0.025***
		(0.002)	(0.003)	(0.003)	(0.003)
	35-40	0.031***	0.016***	0.034***	0.016***
		(0.003)	(0.003)	(0.003)	(0.003)
	40-45	0.036***	0.022***	0.037***	0.023***
		(0.003)	(0.003)	(0.003)	(0.003)
	45-50	0.037***	0.030***	0.037***	0.033***
		(0.003)	(0.003)	(0.003)	(0.003)
	50-55	0.035***	0.030***	0.034***	0.032***
		(0.003)	(0.003)	(0.003)	(0.003)
	55+	0.034***	0.024***	0.033***	0.024***
		(0.003)	(0.004)	(0.003)	(0.003)
	R^2	0.954	0.938	0.104	0.067
	Under-id	-	-	222.451	73.525
	Weak IV	-	-	190.079	49.045

Table 5 cont.					
Tenure	*<1 year*	0.005	0.004	0.000	0.001
		(0.004)	(0.004)	(0.005)	(0.004)
	1-2 years	0.011***	0.008***	0.011***	0.007**
		(0.003)	(0.003)	(0.003)	(0.003)
	2-3 years	0.019***	0.014***	0.020***	0.015***
		(0.002)	(0.003)	(0.002)	(0.002)
	3+ years	0.037***	0.028***	0.041***	0.032***
		(0.003)	(0.003)	(0.003)	(0.003)
	R^2	0.954	0.938	0.106	0.069
	Under-id	-	-	123.684	62.721
	Weak IV	-	-	384.486	255.800
N		4,389,400	4,389,400	2,382,800	2,382,800
N individuals		701,000	701,000	25,200	25,200
N firms		28,539	28,539	406,700	406,700

Notes: See notes to Table 3.

Table 6: Heterogenous rent sharing by firm characteristics

		1	2	3	4
		OLS	IV	Male IV	Female IV
Firm Size	*5-20 FTE*	0.035***	0.028***	0.031***	0.021***
		(0.001)	(0.004)	(0.004)	(0.004)
	20-50 FTE	0.034***	0.033***	0.035***	0.028***
		(0.001)	(0.003)	(0.003)	(0.003)
	50-100 FTE	0.029***	0.033***	0.035***	0.025***
		(0.002)	(0.003)	(0.003)	(0.003)
	100+FTE	0.024***	0.031***	0.033***	0.026***
		(0.003)	(0.003)	(0.003)	(0.003)
	R^2	0.952	0.089	0.103	0.067
	Under-id	-	904	934	594
	Weak IV	-	731	754	536
Firm Age	*0-2 years*	0.021***	0.032***	0.035***	0.025***
		(0.007)	(0.007)	(0.008)	(0.008)
	2-5 years	0.016***	0.021***	0.024***	0.015***
		(0.003)	(0.004)	(0.005)	(0.004)
	5+ years	0.029***	0.032***	0.034***	0.026***
		(0.002)	(0.003)	(0.003)	(0.000)
	R^2	0.952	0.090	0.104	0.067
	Under-id	-	56	124	20
	Weak IV	-	66	178	26
	N	6,772,200	6,772,200	4,389,400	2,382,800
	N firm	29,349	29,349	28,539	25,200
	N indiv	1,107,800	1,107,800	701,000	406,700

Notes: See notes to Table 3

Appendix A: Title of appendix

Table A 1: Estimating the rent thresholds

Year	Proportion of FTE in firms with no excess rent	Pass-through parameter γ
2002	29.1%	0.083***
2003	15.5%	0.085***
2004	22.9%	0.086***
2005	16.0%	0.086***
2006	12.4%	0.083***
2007	34.4%	0.098***
2008	27.9%	0.092***
2009	29.0%	0.091***
2010	21.9%	0.093***
2011	17.1%	0.101***
2012	30.6%	0.125***
2013	21.2%	0.112***
2014	27.4%	0.116***
2015	25.5%	0.116***
2016	32.0%	0.134***
2017	20.9%	0.120***
2018	20.3%	0.130***

Table A 2: Summary statistics for excluded workers and firms

	Zero excess rents	Insufficient wage information
Real annual wage	$54,700	-
	($29,900)	-
Log real annual wage	10.826	-
	(.376)	-
Real quasi-rents pw	$11,200	$32,400
	($5,700)	(173100)
Log real quasi-rents pw	9.093	9.95
	(.901)	(1.113)
Real excess rents	-	$17,100
	-	(172700)
Excess log rents	-	.49
	-	(.705)
Employment (FTE)	769	781
	(1652)	(1865)
Log employment (FTE)	4.805	4.817
	(1.987)	(1.94)
Firm age	23	20
	(19)	(17)
Female	.381	.52
	(.486)	(.5)
European	.675	.625
	(.468)	(.484)
Māori	.157	.184
	(.364)	(.387)
Pacific	.103	.093
	(.304)	(.29)
Asian	.129	.123
	(.335)	(.329)
MELAA	.016	.025
	(.125)	(.157)
Missing ethnicity information	.016	.054
	(.124)	(.227)
Age 25-30	.175	.27
	(.38)	(.444)
Age 30-35	.147	.17
	(.354)	(.376)
Age 35-40	.135	.132
	(.342)	(.339)
Age 40-45	.136	.116
	(.343)	(.321)
Age 45-50	.13	.097
	(.336)	(.296)
Age 50-55	.112	.078
	(.315)	(.268)
Age 55+	.165	.136
	(.371)	(.343)
No qualifications	.178	.158
	(.383)	(.365)
High-school qualifications	.328	.289
	(.47)	(.453)
Post-school qualifications	.221	.169
	(.415)	(.374)

Table A2 cont.

Bachelor's	.084	.096
	(.277)	(.294)
Postgraduate qualifications	.043	.054
	(.203)	(.226)
Missing qualifications	.145	.234
	(.352)	(.424)
Tenure < 1 year	.092	.363
	(.289)	(.481)
Tenure 1-2 years	.227	.399
	(.419)	(.49)
Tenure 2-3 years	.217	.142
	(.413)	(.349)
Tenure 3+ years	.463	.097
	(.499)	(.296)
N	2,213,500	3,244,700
N individuals	923,900	1,389,500
N Firms	25,968	46,278
Avg. obs per job	1.8	1.3

Notes: See notes to Table 1.

Figure A 1: Percent of employment and firms earning zero rents by industry

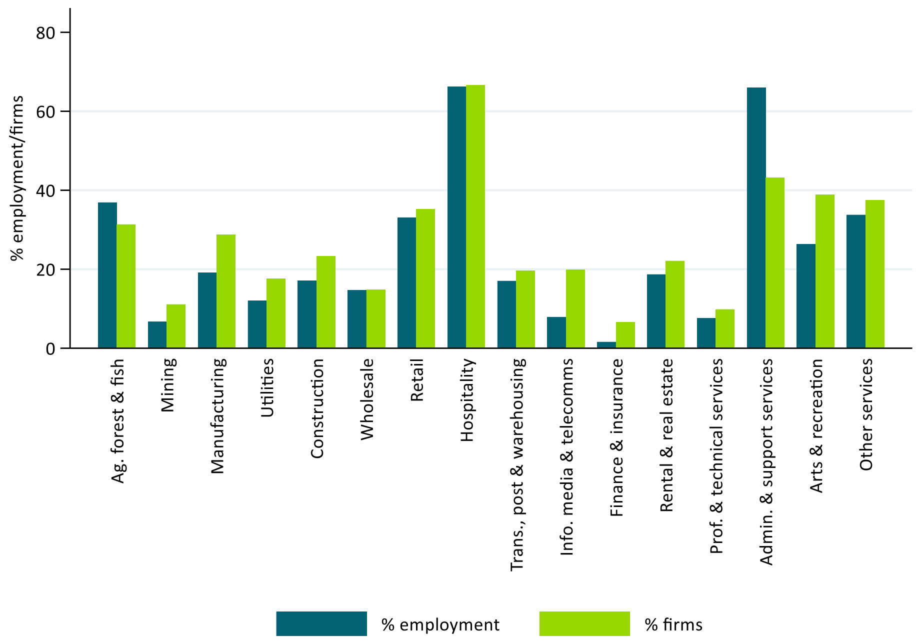


Figure A 2: Estimated log tau over time, 2018 NZD