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Do workers share in firm success?

Pass-through estimates for New Zealand

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These results are not official statistics. They have been created for research purposes from the Integrated Data Infrastructure (IDI) and Longitudinal Business Database (LBD) which are carefully managed by Stats NZ. For more information about the IDI and LBD please visit https://www.stats.govt.nz/integrated-data/.

The results are based in part on tax data supplied by Inland Revenue to Stats NZ under the Tax Administration Act 1994 for statistical purposes. Any discussion of data limitations or weaknesses is in the context of using the IDI for statistical purposes, and is not related to the data's ability to support Inland Revenue's core operational requirements.

Abstract

We study the extent to which firm financial performance is passed on to workers in the form of higher wages and the degree to which this pass-through has changed over the period 2002-2018. We use both value added per worker and a measure of quasi-rents as measures of financial performance. Value added per worker is the standard measure used internationally. Quasi-rents better approximate the resources available to be shared between workers and firms as it takes into account the rental cost of capital as well as the reservation wages of workers. We estimate the reservation wage bill for each firm using estimates from a two-way fixed-effect model. We estimate models similar to those typically used in the international literature and further decompose the estimated pass-through into the contribution from worker sorting and the contribution from rent-sharing. Our instrumental variables estimates of pass-through are in the range of 0.12 and 0.19 for value added and 0.11 and 0.07 for quasi-rents. Worker sorting explains between 35% and 50% of pass-through. While the extent of overall pass-through is relatively stable over time, the contribution of worker sorting declines dramatically to explain almost none of the estimated pass-through. We contribute to the literature by demonstrating a method to calculate quasi-rents, by testing for changes over time in pass-through, and examining the relative importance of worker sorting over time.

JEL classification

J31 Wage Level and Structure - Wage Differentials;

J71 Discrimination;

E25 Aggregate Factor Income Distribution

D22 Firm Behavior: Empirical Analysis

Keywords

Wage determination; Rent sharing; Worker sorting

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1 Introduction

Almost all advanced economies have experienced sluggish wage growth in the years following the global financial crisis (GFC) in 2008/2009, despite strong growth in employment and declining unemployment rates (Blanchflower 2019). This, coupled with the long-run declines in their labour income shares (LIS), raises significant concerns about inequality, the balance of power between workers and owners of capital, and whether the benefits of economic growth are being widely shared (e.g. OECD 2015; Piketty 2014; Stansbury & Summers 2020).

New Zealand is no exception, having seen a long-run decline in the LIS and relatively low nominal wage growth post GFC (e.g. Rosenberg 2010, 2017; Conway et al. 2015; Fraser 2018; Hyslop & Rice 2019; Ball et al. 2019). These developments raise questions around whether workers are getting a fair share of the recent economic success. The issues of inclusive growth and worker power are live policy issues in New Zealand. There is current policy work underway on 'Fair Pay Agreements', which provide a framework for bargaining in sections of the labour market where bargainning is currently difficult (Fair Pay Agreement Working Group 2018).

We contribute to this discussion by examining inclusive growth at the firm level. We examine the extent to which firm performance is reflected in average wages ("pass-through"). We consider two main measures of firm performance. The first is value added per worker, a standard measure of labour productivity and one that has been used in most international studies on this topic (see Card et al. 2018). The second is a measure of 'quasi-rents', which more closely approximates the amount that is available to be shared as wage premiums above reservation wages. Unlike value added, quasi-rents account for the cost of capital and reservation wages of workers. We make use of estimates from a two-way fixed effect model, of the form introduced by Abowd et al. (1999), to estimate individual reservation wages and firm reservation wage bills.

Our aim in this paper is to explore pass-through in the private market sector of the economy. We document both the overall cross-firm and within-firm patterns of pass-through. We further decompose the overall pass-through estimates into the contribution of worker sorting and the contribution of rent-sharing, following the method of Card et al. (2018). Worker sorting is the concentration of highly-skilled workers (who attract higher wages) in better performing firms who are able to pay them higher wages. Rent-sharing occurs when firms share (and/or workers demand) a share of the economic rents earned by the firm. We then look at whether pass-through and the contributions of worker sorting and rent-sharing have changed over the period 2002-2018.

OLS estimates of the pass-through elasticity range between 0.03 and 0.13 for value added per worker, and 0.01 and 0.05 for quasi-rents per worker. The variation depends on the stringency of other controls included in the model. Instrumental variables (IV) estimates are larger, between 0.12 and 0.19 for value added, and 0.07 and 0.11 for quasi-rents. While the pass-through elasticity for quasi-rents is lower than for value added (as quasi-rents are smaller than value added), it implies a larger dollar value of pass-through for a given dollar increase in quasi-rents or value added. In our preferred IV specification, workers receive an extra 8c for every extra dollar of quasi-rents, and 5c per dollar of value added. Overall, our estimates are consistent with those from international studies.

We find that worker sorting explains approximately 40%- 50% of the pass-through for value added and 30%-40% for quasi-rents. This is expected as quasi-rents takes worker quality into account by subtracting the reservation wage-bill. Our results for the contribution of worker sorting to overall pass-through are similar to those in Card et al. (2018) for Portugal.

Our estimates of overall pass-through display a pro-cyclical pattern, with estimates 1-2 percentage points lower in 2008-2010 during the GFC. Estimates recovered from 2011 onwards to approximately the same level as in the pre-2008 period. Our decomposition results show a marked change in the importance of worker sorting over the period. Worker sorting explained around 60% of value-added pass-through and 40%-50% of quasi-rents pass-through over the period 2002-2007. The sorting contribution falls virtually to zero in the latter half of the sample period. In some specifications, the contribution of worker sorting to pass-through becomes negative.

The most closely related New Zealand microeconometric study into questions of rent-sharing is Sin et al. (2020). The authors consider the role of differential worker sorting and rent-sharing in explaining the gender wage gap in New Zealand. Our interest in this paper is documenting overall patterns in pass-through to average firm wages, patterns in worker sorting and rent-sharing. We also consider how these overall patterns have changed over time. New Zealand is an interesting case study for these issues as it hasn't experienced the same rise in wage inequality or the same increase in the dispersion of firm performance since 2000 that other advanced economies have (e.g. Berlingieri et al. 2017; Criscuolo et al. 2020). We also demonstrate a way to calculate quasi-rents, which better approximates the surplus available to be shared in the form of higher wages or taken as profits. We consider only the sharing of rents as wages and do not examine other possible uses of rents by the firm. Firms could use a portion of these rents to fund expansion, fund R&D, to provide a cash buffer for protection against negative shocks, or to pay dividends to firm owners.

Our results show that yes, workers do share in firm success. Most of this is due to better performing firms paying higher wages in general, although there is some pass-through of improvements in performance as higher wages. The firms people choose to work at does matter for their wages, but the general macroeconomic conditions exert a significant influence on wage growth.

The paper is structured as follows. Section 2 provides a literature review and an overview of the New Zealand labour market and institutional context. Section 3 presents our analytical framework. Section 4 describes the data used in this paper. Section 5 presents our results, and section 6 concludes.

2 Background and literature review

In order to understand the process of how firm performance is passed through to wages, we need to look beyond simple models of the wage determination, and consider how rents are generated and shared in imperfectly competitive labour markets. In simple competitive models of the labour market, wages are determined as the value of the marginal product of labour (*p.MPL*) – the change in output that results from an increase in labour, valued at the price at which output is sold. When MPL increases by 10%, the wage also increases by 10% - a pass-through of 100%. Our study of pass-through focuses on the relationship between average labour productivity (output per worker) and wage levels. This may differ from 100% due to the degree of substitutability of labour with other inputs of production, or due to imperfect competition, either of which results in rents that are shared unequally between workers and owners of firms.

Maintaining marginal productivity as a fixed proportion of average productivity when output changes requires a very specific type of production function (e.g. Cobb-Douglas). More generally, the relationship depends on how substitutable labour is with other inputs. Strong substitutability implies that firms reduce employment proportionally more than the wage increases, resulting in a lower share of revenue going to labour as production is increased. However, even for strong plausible values for substitutability, the implied pass-through is still relatively strong.[1]

More generally, the pass-through of labour productivity to wages can reflect the extent to which economic rents are shared between firms and workers. There are two main sources of rents that are relevant for the analysis of productivity pass-through – product market rents and labour market rents.

Product market rents may arise from imperfect product market competition, whereby firms face a downward sloping demand curve for output, and have some (monopoly) power to set output prices at a mark-up over cost. The resulting economic rents may be sustained by barriers to the entry of competing firms, or they may be 'quasi-rents', defined as rents that are temporary and will eventually be eroded by the entry of competitors or the loss of intellectual property protections. Monopolistic firms sell a lower-than competitive level of output at a higher-than-competitive price, and make a positive profit. If they operate in a competitive labour market, they would employ fewer workers due to the reduced level of output, but would pay above-market wages only if they were to share their profits. Whether changes in firm performance are reflected in wage levels would depend on how responsive this sharing is when rents change.

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[1] For example, If the elasticity of substitution in a two-factor CES production function were 2 (twice as high as for the Cobb-Douglas production function), the pass-through of labour productivity changes to wages falls to 50%. A pass-through of 10% would require an unrealistically high elasticity of substitution of 10.

In contrast to product market rents, labour market rents arise where firms can pay wages at a level below the marginal value of what workers contribute to output. In imperfectly competitive labour markets, firms may have to raise wages if they want to increase employment, and they may have some (monopsony) market power to set wages as well as employment. This enables them to pay wages below the marginal value of what workers produce and they can thus earn rents that could potentially be shared with workers. Labour market rents may also arise from adjustment costs in the labour market. If it is costly for workers to search for new jobs, and costly for firms to recruit new workers, both workers and firms gain from keeping a job going and this (dynamic monopsony) is a source of rent that can potentially be shared. When firm performance increases, monopsonistic firms will increase employment, and wages may rise not only because they are unable to expand employment without raising wages but also because the amount of rents shared as wages may change.

In this study, we do not attempt to identify the various potential sources of rents, focusing instead on the degree to which any rents that do exist are shared. The way that rents are shared is determined separately from production or pricing decisions. The share of rents that go to workers will therefore reflect relative bargaining power of workers, and may be affected by subjective factors such as what employees or firm owners perceive as fair. The overall measure of pass-through that we aim to estimate is thus a composite measure that summarises potentially different degrees of sharing of potentially heterogeneous types of rent.

2.1 Literature review

The long-run decline in the LIS that has occurred in most advanced economies is at odds with the first of the regularities observed by Kaldor (1957). This states that the share of national income going to labour and capital is constant in the long-run, meaning that the long-run elasticity of the average wage with respect to national income per worker is equal to one. A decline in the LIS means that growth in average wages has not kept pace with growth in national income.

A number of studies have sought to explain the decline in the LIS within a perfectly competitive macroeconomic framework. These studies typically consider the impact of significant changes in the macro-environment, such as the increase in import competition (Autor et al. 2016) or the role of technology (Fraser 2018; Autor & Salomons 2018).

Other macro-level studies have examined the role of changing product market competition. A seminal contribution is Autor et al. (2020), which examines the role of 'superstar' firms in explaining the decline in the US LIS. Superstar firms are characterised by very high value added per worker but a relatively low LIS. Autor et al. (2020) find that the decline in the aggregate LIS is driven largely by reallocation of workers and output to these firms, rather than a decline in the LIS for the average firm.

Stansbury & Summers (2020) put forward an alternative hypothesis for the decline in the labour income share – a decline in worker power. Worker power acts as a countervailing force to firm labour market power and means workers are able to capture a portion of the rents earned by firms (and to which the workers contribute). They show for the US that the unionisation rate has declined along with the union wage premium, the firm size premium and industry premiums, as has the relationship between industry productivity and wages. The authors suggest these changes signal a decline in worker power and this can explain not only the decline in the LIS, but also rising profitability and market valuations of businesses, slow wage growth, and a decline in the NAIRU. [2]

A large literature exists which considers the extent to which workers capture a share of the rents earned by firms (see Card et al. 2018). As rents vary across firms, this is a possible explanation as to why wages for otherwise similar workers vary across firms. There are a number of approaches to modelling this relationship. Some rent-sharing studies use a bargaining framework between a firm and a union/worker to understand the links between rents and wages in terms of relative bargaining power (e.g. (Blanchflower et al. 1990; Abowd & Lemieux 1993; Van Reenen 1996; Card et al. 2014). Others use models where firms have either product or labour market power (or both) i.e. face a downward-sloping product demand curve and/or an upward sloping labour supply curve (e.g. Kline et al. 2019;, Card et al. 2018). Others use rent-sharing models to link increases in the dispersion of performance across firms to increases in wage inequality (e.g. Barth et al. 2016, Criscoulo et al. 2021).

Early rent-sharing studies use firm-level data to relate average wages to measures of economic rents. Typical rent measures used in the earlier literature include profits per worker and quasi-rents per worker, where quasi-rents is usually measured as sales per worker less an alternative wage, typically an industry average (e.g. Blanchflower et al. 1990; Abowd & Lemieux 1993; Van Reenen 1996; Hildreth & Oswald 1997; Hildreth 1998).[3] Earlier studies find rent-sharing elasticities between 0.2 and 0.3, depending on the performance measure used. These are likely overestimates of the rent-sharing parameter as these earlier studies are unable to control for worker heterogeneity across firms. If better quality workers, who attract higher wages regardless of where they work, are more likely to work in or sort into high rent firms, this will generate a positive correlation between wages and rents and lead to an overstatement of the rent-sharing elasticity.

More recent studies use linked employer-employee data and are better able to control for worker heterogeneity and sorting (e.g. Card et al. 2014, 2016; Carlsson et al. 2016; Guertzgen 2009; Arai & Heyman 2009; Andrews et al. 2019). Studies of this type often employ a job-stayers design, relating changes in rents to changes in wages for workers

[2] The NAIRU is the non-accelerating inflation rate of unemployment.
[3] Abowd & Lemieux (1993) use the most comprehensive measure of quasi-rents by subtracting capital costs, intermediate costs, and reservation labour costs. Due to data limitations they use industry-level data on capital and intermediates to create their firm-level measure of quasi-rents. Their measure of the alternative wage is the industry average.

who remain at the firm. This removes the influence of worker sorting by fixing the workforce composition within firms. Rent-sharing elasticities from these studies are lower than the earlier estimates from firm-level studies, typically in the range of 0.05-0.15. This shows the importance of controlling for worker heterogeneity and sorting in these models.

Some studies have explored heterogeneity in the relationship between rents and wages. Arai & Heyman (2009) find that there is a positive and stable effect of profits on wages only for firms with increasing profits. Guertzgen (2009) finds that wages are positively related to rents in the non-union sector and under firm-specific union contracts, but the association is significantly lower under industry-wide contracts. Bell et al. (2019), Andrews et al. (2019), and Stansbury & Summers (2020) all find evidence that the rent-sharing elasticity has declined over time. Stansbury & Summers (2020) use industry-level data from the US and show a decline over the period 1958-2011. Bell et al. (2019) show a decline in the rent-sharing elasticity in the UK over the period 1983-2016 and that this decline is more pronounced among firms with greater product-market power. Andrews et al. (2019) find that the relationship between value added per worker and wages in Australia is significantly lower in the 2013-2016 period than in the 2001-2012 period.

A related literature, which uses similar empirical models, considers the role firms play in providing a type of income insurance for workers. The key feature of insurance models is that firms isolate their workers from temporary swings in firm performance by providing wage stability. The presence of insurance-type behaviour will lower the association between firm performance and wages. Guiso et al. (2005) find that firms fully insure workers against temporary output shocks, while providing partial insurance against permanent shocks. Cardoso & Portela (2009) find similar results and show that collective bargaining and minimum wages are associated with more insurance, and that workers receive more protection against permanent shocks than managers.[4] Juhn et al. (2018) find that insurance is strongest against temporary shocks to revenues and that insurance is weakest for employees in professional services who are in the top 5% of their employers' earnings distribution.

Our work is also related to the literature which considers the role of worker heterogeneity, firm heterogeneity, and worker sorting in overall wage inequality. A common empirical model used in this literature is the two-way fixed effect model, introduced by Abowd et al. (1999). This model relates wages to a set of time-varying individual characteristics, a transferable individual wage premium, a firm wage premium, and an idiosyncratic component which is often interpreted as a worker-firm match premium. These studies typically find that cross-sectional variation in the firm wage premiums explains a substantial portion of the overall wage variation, in the order of 15%-30% (e.g. Abowd et al. 1999; Maré & Hyslop 2006; Card et al. 2013, 2018; Andrews et al. 2008, 2012; Song et al. 2019). Card et al. (2013) and Song et al. (2019) use two-way fixed effect models to consider the role of worker heterogeneity, firm heterogeneity, and worker sorting in explaining trends in wage inequality in West

[4] This could also be interpreted as managers capturing a larger share of the rents.

Germany and the US, respectively. Both find evidence that the variance of firm wage premiums has increased and that the tendency for highly-skilled workers to sort into high premium firms has also increased.

The work of Card et al (2016, 2018) discusses the relationship between the rent-sharing and two-way fixed effect literatures. Card et al. (2016) estimate gender-specific firm wage premiums and decompose the gender wage gap into a sorting component and a bargaining component. They show that women are more likely to work in firms with lower pay premiums and that women receive a lower share of firm-specific rents than men.[5] Card et al. (2018) expand on the earlier work and make the observation that if the two-way fixed effects model is an appropriate specification for wages, then cross-sectional estimates of pass-through elasticities can be decomposed into a worker-sorting component and a rent-sharing component. This is achieved by taking the components of a two-way fixed effects model (covariate index, worker fixed effect, firm fixed effect) and relating each of the components to a measure of rents. Card et al. (2018) find that worker sorting explains 40%-45% of the estimated pass-through elasticity. They also find that the rent-sharing estimate is similar between less educated and more educated workers, but worker sorting is more important for more educated workers.

2.2 The New Zealand labour market and institutional context

The prevalence of rents and the way that they are distributed vary over time and across labour markets, reflecting institutional differences and changes, aggregate income variation, and cyclical variation. In this section, we outline some key features of the New Zealand labour market and recent labour market change. We provide an overall summary of trends in output, GDP and wage growth, and discuss some institutional features of the New Zealand labour market that influence wage setting behaviour.

Our study covers the period 2002 to 2018, a period of moderately strong employment growth, apart from a GFC-related contraction in 2008-09 (Maré 2017). Although the longer term trend in the labour income share in New Zealand has been one of decline since the 1970s, the share has risen since 2002, from 49.5% to 53.8% in 2019. There was a pronounced fall in the labour income share in the late 1980s, at a time when New Zealand undertook widespread reform, including deregulation, privatisation, and state sector reform (Bridgman & Greenaway-McGrevy 2018). It declined further in the 1990s, reflecting ongoing changes in the labour market and labour market institutions (Rosenberg 2017; Conway et al. 2015).

As noted in the introduction, the labour income share, low wage growth and concerns about whether the economic growth is being shared equitably are live issues in New

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[5] Sin et al. (2021) apply the methods of Card et al. (2016) in their study of the determinants of the gender wage gap in New Zealand. They find that both differential worker sorting and rent-sharing contribute to the gender wage gap. Women are more likely to work in firms with less rents to share. Women also receive a smaller portion of rents than men regardless of the level of rent available.

Zealand. Like many other advanced economies, New Zealand has also had a prolonged period of low productivity growth (Conway 2016; 2018).

A comparison of recent trends in New Zealand's labour market compared with the rest of the OECD is shown in Table 1. The New Zealand labour market can be characterised as having high rates of employment and participation but relatively low wages and labour productivity. New Zealand's employment and participation rates are relatively high compared to the OECD average and have been growing more rapidly over the last two decades, even with relatively strong population growth driven by net migration (Maré 2018). The unemployment rate has generally been lower in New Zealand. The workforce has similar levels of tertiary qualifications to the rest of the OECD.

Table 1: Key facts about the New Zealand labour market

	New Zealand		OECD	
	Latest available	Avg. annual growth, 2000-latest	Latest available	Avg. annual growth, 2000-latest
Employment growth	-	2.0%	-	1.2%
Employment rate	77.4%	0.37	68.8%	0.17
Participation rate	80.9%	0.31	72.8%	0.15
Unemployment rate	4.3%	-0.1	5.6%	-0.04
Median real wage	$44,000	2.0%	$48,600	1.0%
Min wage as % of median	65.9%	0.82	54.2%	0.59
GDP per hour+	$42	1.0%	$54	1.2%
GDP per capita+	$39,400	1.61%	$42,500	1.2%
LIS+	48.9%	0.14	54%	-0.1
50/10 ratio+	1.47	-0.01	1.65	0
90/10 ratio+	2.71	0	3.31	-0.01
90/50 ratio+	1.84	0.01	2.01	0
% 25-64 with tertiary qual	39.1%	-	38.0%	-
Trade union density	17.3%	-0.24	24.8%	-0.61

Notes: + denotes the latest available data is for 2018, otherwise 2019 data are used. Growth in ratios and percentages is a percentage point change. The median real wage is measured in constant 2019 USD. GDP per hour and GDP per capita are measured in constant 2015 USD. Source for all data is OECD.stat. The labour income share is measured as compensation of employees over compensation of employees plus gross operating surplus and gross mixed income.

While employment and participation rates are high, wages in New Zealand are low compared to the OECD average. The median full-time annual wage was 90% of the OECD average in 2019. New Zealand's minimum wage is also relatively high, sitting at 66% of the median wage (compared to 54% in the OECD) and the minimum wage is scheduled to increase further in April 2021. The New Zealand wage distribution is typically narrower at all parts of the distribution, although the gap between the top 10% and the median has been increasing slightly. The bottom half of the distribution has become more compressed, likely a result of the increases in the minimum wage. The adult minimum wage has increased by 75% in real terms since 1999 (Maré &

Hyslop 2021). A high minimum wage is likely to lead to a lower pass-through estimate as shown by Cardoso & Portela (2009).

New Zealand labour productivity, measured as GDP per hour worked, sits around 78% of the OECD average. This is consistent with New Zealand's low wages compared to the rest of the OECD. GDP per capita sits at 93% of the OECD average, reflecting New Zealand's higher employment rate.

Figure 1: The labour income share in New Zealand and the OECD, 1990-2019

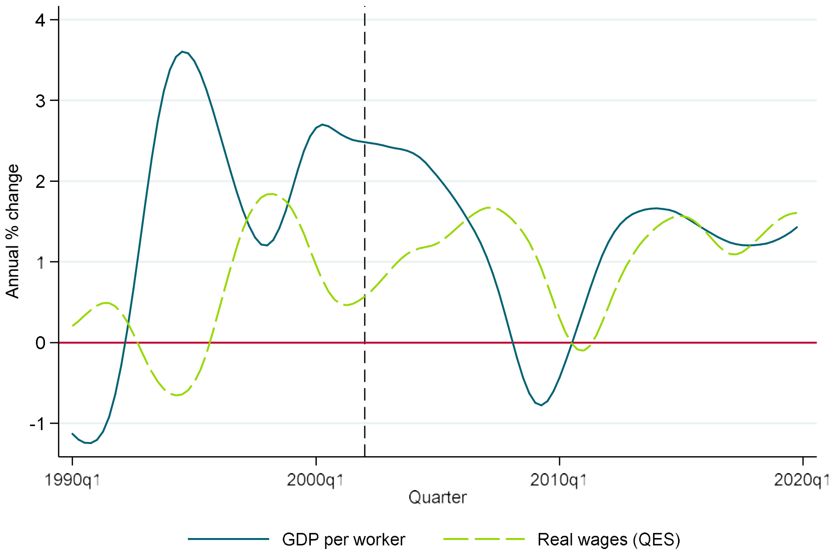
Source: OECD.stat

Like many other countries, New Zealand has seen a long-run decline in the labour income share. Figure 1 plots the labour income share in New Zealand and the OECD over the period 1990-2018. The LIS declined over the 1990s, continuing a trend that began in the 1980s (Bridgman & Greenaway-McGrevy 2018; Rosenberg, 2017). However, the labour income share began increasing from 2000, rising from 45% to 50% in 2009. From this point, the LIS has remained in the range of 48%-50%. In contrast, the trend in the OECD average is one of decline. In 1990, the LIS was 56%, declining to 54% in 2018. This shows that growth in labour productivity has exceeded wage growth across the OECD. For New Zealand, there have been periods where wages have grown more rapidly than labour productivity and vice versa. However, the long-run trend in New Zealand's labour income share is one of decline. Rosenberg (2010) shows that, over the period 1978-2006, average wages grew by 44% while labour productivity grew by 90%, giving an aggregate pass-through elasticity of 0.49 over the period.

The period from 2002-2018 is our study period. Over this more recent period, wages have generally grown more strongly than value added per worker. Our study period misses out the large declines in the labour income share in the 1980s and 1990s that coincided with the period of large structural reforms to the New Zealand economy.

Figure 2 more directly compares growth in labour productivity (measured as GDP per worker) against growth in average wages over the period 1990-2020. During the first part of the period, growth in GDP per worker was above growth in real wages, consistent with a declining LIS. The correlation between the two is also relatively weak during the period 1990-2010. Growth in real wages began to accelerate in the early 2000s, as growth in GDP per worker began to decline. This is around the start of our sample period. Both saw large decreases in growth during the GFC. Post GFC, both have averaged approximately 1.1% growth.

Figure 2: Growth in GDP per worker and real wages (CPI adjusted)[6]



Source: StatsNZ

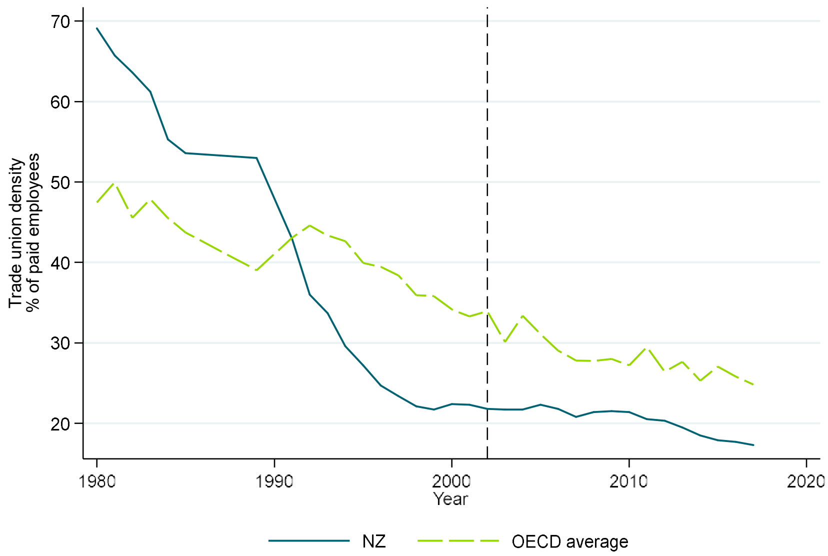
Another major change in New Zealand's labour market over the last 40 years has been the declining rate of union membership. Figure 3 plots the trade-union density in New Zealand and the OECD average from 1980-2017. While the trade union density has declined in New Zealand and the OECD in general, the fall in New Zealand was much more dramatic. New Zealand went from one of the most unionised OECD countries to one of the least. Over 50% of paid employees were union members during the 1980s. This declined rapidly from 1990 onwards, levelling off at around 22-24% in the late 1990s. From there it has declined slowly to sit around 17% in 2017, compared to 25% in the OECD. Union density is highest in the public sector, with 60% of employees being members of a union, compared to 10% in the private sector. A similar pattern is seen in other countries (Ryall & Blumenfeld 2019).

The primary legislative framework that governs the relationship between employees and employers in New Zealand is the Employment Relations Act 2000 (ERA). This replaced the earlier Employment Contracts Act 1991, which deregulated employment

[6] The plotted series are the trend estimates from an HP filter with the value of the smoothing parameter set at 200.

relationships and reduced the role of unions to bargaining agents with the same status as other actors in the market (Rosenberg 2017). The ERA restored the special collective bargaining rights of unions.

Figure 3: Trade union density in New Zealand and the OECD, 1980-2017



Source: Ryall and Blumenfeld (2017); OECD.stat

Individual employee-employer bargaining is the predominant form of bargaining in New Zealand. Where collective bargaining occurs in the private sector, agreements are almost universally between a union and a single employer. While the New Zealand legal framework does allow for multi-employer bargaining, it is not common (Rosenberg 2017). The results of Guertzgen (2009) suggest that individual-level bargaining and firm-specific collective agreements result in similar rent-sharing elasticities, while multi-employer bargaining results in lower rent-sharing. This finding could also be interpreted as multi-employer bargaining delivering a greater degree of income insurance to workers, consistent with Cardoso & Portela (2009).

Changes to collective bargaining arrangements are being actively considered in New Zealand. The most recent work in this area is the development of proposed Fair Pay Agreements. The proposed agreements allow for sector-wide bargaining in sectors or occupations that meet certain criteria. The agreements will be given legal effect by government and apply to all employers and employees in that sector/occupation. A Fair Pay Agreements Working Group was established in 2018 and delivered a set of recommendations (Fair Pay Agreement Working Group 2018). Consultation on the proposed agreements was conducted in 2019 (Ministry of Business, Innovation and Employment 2019).

3 Data and summary statistics

3.1 Data

Our data are drawn from the Longitudinal Business Database (LBD) and Integrated Data Infrastructure (IDI).[7] These databases contain a rich set of survey and administrative data on businesses (LBD) and individuals (IDI), and are both managed by StatsNZ. Our population of interest is private-for-profit employing firms in the measured sector, along with the employees in these firms.[8] Our sample of firms comes from the LBD productivity tables described in Fabling & Maré (2015; 2019). These tables combine information from the Annual Enterprise Survey and IR10 financial summary forms to produce a harmonised set of annual firm-level financial information for use in micro-econometric analysis. From this table we take information on gross output, intermediate expenditure, and the cost of capital.

We then identify the population of workers at firms in our sample using the IDI and LBD labour tables (Fabling & Maré 2015). This table contains monthly job-level information for all employees in New Zealand and is derived from Inland Revenue's (IR) Employee Monthly Schedule (EMS).[9] We take information on monthly earnings, estimated full-time equivalent (FTE) employment, age, and gender. We also take the estimated firm and worker wage premiums derived from a 2-way fixed effect model, similar to Abowd et al. (1999). We aggregate the monthly information to an annual-equivalent basis.

3.1.1 Defining our sample

We restrict attention to firm-year observations with non-zero paid employment (i.e. exclude working proprietor only firm-year observations). We further restrict our sample to those firms with non-missing data on gross output, intermediate expenditure, and capital. Our sample of firms is an unbalanced panel containing information on 319,299 firms over the period 2002-2018 for a total of 1,713,759 observations.[10] This is equal to 71% of firm-year observations and 81% of employment in the private-for-profit measured sector. For our analysis sample, we restrict attention to firms with at least one employee with one month of usable earnings data (i.e. drop firms with adjusted annual FTE input < 1/12). This leaves us with a final analysis sample of 1,409,835 observations on 266,580 firms.

[7] For more information on the LBD, see Fabling & Sanderson (2016).

[8] The measured sector is defined as "industries that mainly contain enterprises that are market producers. This means they sell their products for economically significant prices that affect the quantity that consumers are willing to purchase" (Statistics New Zealand 2014). In practice this excludes the Public Administration and Safety, Education and Training, and Healthcare and Social Assistance industries.

[9] The EMS are the monthly payroll returns that firms file with Inland Revenue for the purposes of administering New Zealand's PAYE income tax system.

[10] All of our data come from the January 2020 IDI Refresh (IDI_Clean_20200120) and the December 2019 LBD archive (ibuldd_clean_december_2019).

3.1.2 Measuring wages

We use annual FTE earnings as our wage measure as we do not directly observe hourly wages or hours worked in our data. In constructing this measure, we place several restrictions on the job months that we include as some job-months provide better information about the underlying wage rate than others.

We exclude the first and last months of an individual's employment at a firm from the earnings calculations. Earnings in the first and last months of employment are unlikely to accurately reflect a person's regular earnings due to starting or leaving part way through a reporting month or payments associated with starting or leaving a job (signing bonus, pay-out of annual leave etc.). We further exclude job-months where the employee is obviously part-time.[11] Including start and end months and months where the minimum wage is binding would lead us to underestimate the underlying wage rate for some workers and firms. We calculate the firm-level average wage by taking the sum of earnings in included job-months for the year and dividing by the annual included FTE employment.[12]

Figure A1-A3 in Appendix A show the importance of these restrictions. Figure A1 shows the percentage of job-months excluded both overall and separately for spell starts and ends and obviously part-time months. The part-time constraint is the main reason why job-months are excluded and this has become more important over time.[13] We exclude approximately 1/3 of job-months, but these job-months account for 15%-20% of FTE employment and 10%-12% of total wages.

Excluded job-months are a non-random selection. They are generally a combination of part-time, low pay, or temporary/short term jobs. People working in these types of jobs tend to be younger, are more likely to be women, and are more likely to be migrants. These types of working arrangements are also more prevalent in certain industries, such as agriculture, retail trade, and hospitality.

Figure A3 shows the cross-industry variation in the fraction of earnings and FTE we include. A relatively high proportion of FTE and earnings are excluded in the retail, hospitality, and administrative and support services industries, while retention rates are highest in finance and insurance, information media and telecommunications, mining, and manufacturing. Industries with a large fraction of earnings excluded have a lot of

[11] 'Obviously' part time is when the total monthly earnings is less than what an individual would earn working 40 hours per week at the minimum wage. Estimated FTE for these jobs is then the ratio of observed earnings to full-time minimum wage earnings (see Fabling & Maré 2015). FTE-adjusted earnings for these workers are then simply the minimum wage. This will overstate the proportion of workers at the minimum wage and therefore understate the variance in wages. Some highly-paid part time workers will be included although we will underestimate their wages. We do not have sufficient information to distinguish between highly-paid part-time workers and low-paid full time workers where monthly earnings exceed full-time minimum wage earnings.

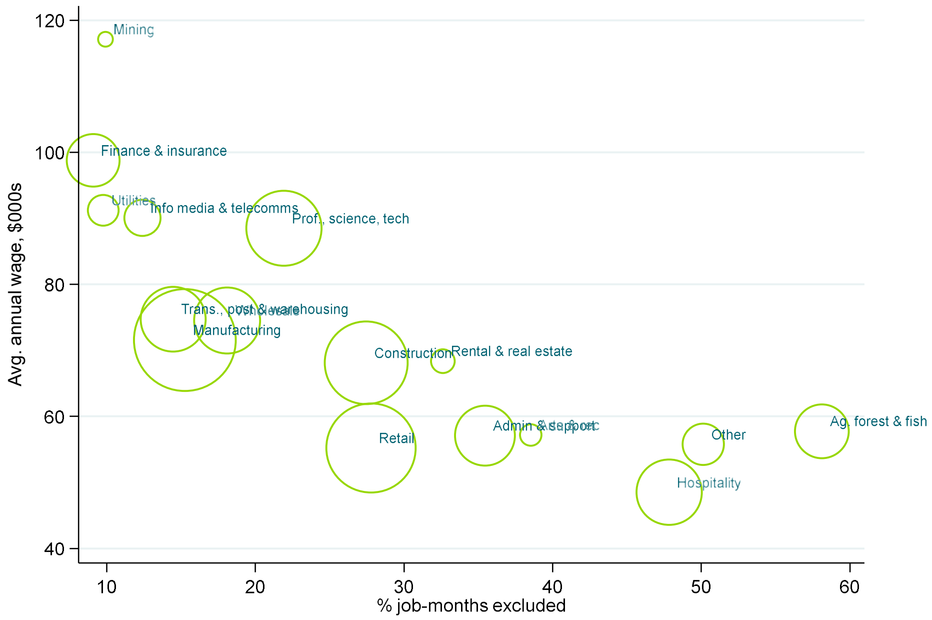
[12] We assume that the wage rates of included workers are representative of those we exclude.

[13] One reason for this is the gradual decline in average hours worked along with the large increase in the minimum wage (relative to the median) over the period.

part-time or casual workers and have relatively high rates of worker turnover. They also tend to be relatively low wage, as shown in Figure 4. This plots the fraction of job-months excluded over the entire 2002-2018 period against our preferred measure of wages.[14] There is a clear negative relationship between the fraction of job-months excluded and our wage measure.[15] To account for this in our analysis we construct the fraction of FTE included in the wage calculation to total FTE to include as a control variable in our regressions.

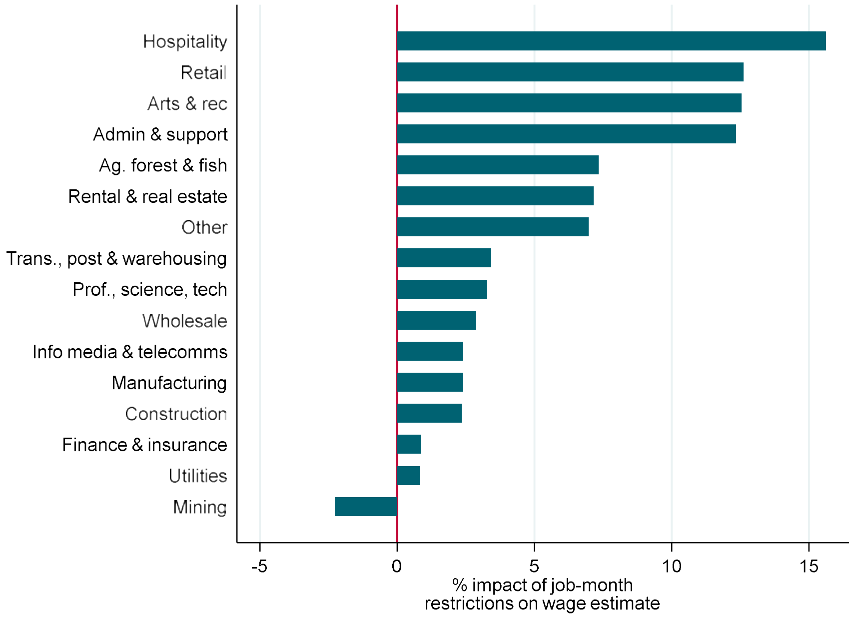
The industries where we exclude most job-months are also the ones where the difference between a simple wage estimate (firm wage bill over firm FTE) and our preferred estimate is largest. This is shown in Figure 5. Our measure of wages is over 15% greater than a simple estimate in the hospitality sector, and around 12% higher in retail, arts and recreational services, and administrative and support services.

Figure 4: Our wage estimate and the proportion of job-months excluded by industry



[14] The size of the bubble represents the share of employment in each industry.
[15] This pattern also holds when using an unadjusted wage measure.

Figure 5: Percent difference between our preferred wage measure and an unadjusted wage measure (wage bill/FTE)[16]



3.1.3 Measuring firm performance

Our key variable of interest is a measure of the economic rents available to be split between workers in the form of higher wages or firm profits. We calculate two measures of rents: value added per worker and quasi-rents per worker.

Value added per worker is the standard measure used in the international literature (see Card et al. 2018). We calculate value added as $VA = GO - M$ where VA is value added, GO is gross output, and M is intermediate expenditure.[17] We calculate value added per worker by dividing this by the firm's FTE labour input.[18] Value added per worker represents the resources left over after paying material costs.

Our second measure of rents is quasi-rents per worker. Following Card et al. (2018), we define quasi-rents as:

$$QR = VA - (r + \delta)K - bL$$

[16] Our preferred wage measure is lower than a simple wage bill/FTE measure for the mining sector. The most likely explanation for this is the exclusion of short-term jobs with extremely high wage rates.

[17] M excludes any rental and leasing of capital goods. These costs are included in the measure of capital.

[18] We exclude the labour input of any working proprietor. We don't have labour income data for working proprietors. Excluding working proprietor input will overstate labour productivity but, given our interest is in how any surplus is split between workers and firms, part of the surplus not shared with workers will be taken as income for any working proprietors.

where QR is quasi-rents, VA is value added (defined above), $(r+\delta)K$ is the cost of capital plus depreciation, L is labour input and b is the alternative wage.[19] Quasi-rents can be thought of as the profit a firm would earn if workers are paid their reservation wage.

A key issue in calculating quasi-rents is getting a measure of the reservation wage as this is not directly observable. Other papers that have constructed a measure of quasi-rents typically use an industry-average wage as the measure of b (Abowd & Lemieux 1993; Van Reenen 1996; Guertzgen 2009). There are a number of issues with using an industry average as a measure of the reservation wage. First, a high proportion of firms in the industry will be paying wages below this measure of the reservation wage. Second, it assumes that all workers in the industry have the same reservation wage.

To estimate b, we make use of the estimated individual and firm earnings premiums estimated from a two-way fixed effect model to calculate a worker-specific reservation wage. The estimates we use are derived from the equation:

$$\ln w_{ijt} = a_i + \phi_j + Z_{ijt}\lambda + \tau_t + \xi_{ijt} \tag{1}$$

where a_i is the estimated individual earnings premium for worker i, ϕ_j is the estimated firm earnings premium for firm j, $Z_{ijt}\lambda$ is an index of observable time-varying worker characteristics, τ_t are year dummies, and ξ_{ijt} is the error term.[20] Our estimate of b is then:

$$b_{it} = e^{a_i + Z_{ijt}\lambda + \tau_t} \tag{2}$$

We aggregate our individual measure of b to the firm level to calculate the average reservation wage and the reservation wage bill. This measure aims to capture the wage a person would expect to earn at a firm that paid no wage premium (i.e. $\phi_j = 0$). However, the estimated b will incorporate the average level of premiums received by an individual, or the level correlated with worker characteristics. This level may be positive on average.

This means that the normalisation of the firm earnings premium is important.[21] We jointly calibrate the firm earnings premium and the reservation wage to ensure that the reservation wage includes only a minimal level of premium, and the firm premium is non-negative for most firms. The firm effects available in the Fabling & Maré (2015) labour tables are normalised to be (FTE-weighted) mean zero, meaning that firm premiums are measured relative to the firm of the average worker. Using the mean-zero firm premiums has the same problem as using some industry-average wage to estimate b, namely that a large number of workers would appear to be paid below their

[19] The key difference between quasi-rent and profit is in the labour cost term. Profit is defined as $\pi = VA - (r+\delta)K - wL$, where w is the wage rate paid by the firm. In order for a firm to have a positive level of employment, it must pay its workers at least the reservation wage (i.e. $w \geq b$). In this view, profit is the return to the business owner after any portion of rent has been shared with workers.

[20] The variables included in *Z* are gender-specific quartics in age.

[21] The firm and worker fixed effect estimates are not uniquely identified, meaning that some normalisation is required.

reservation wage. To solve this issue, we re-centre the distribution of the (FTE weighted) firm earnings premiums so that $E(\phi_j) > 0$ by adding the value at the first percentile to the estimates provided in the labour tables.[22] This means the firm pay premiums are now measured relative to the first percentile and that 99% of workers will be earning their reservation wage, plus some positive premium.[23]

Our estimate of b has two key advantages over industry averages. First, by construction, the vast majority of workers are paid at least their reservation wage. Second, our estimate of b varies across individuals. We might expect that older workers have higher reservation wages than younger workers, or that more highly-skilled workers have higher reservation wages than lower-skilled workers. Utilising the parameters from the two-way fixed effect model allows us to estimate each individual worker's reservation wage and to estimate each firm's reservation wage bill (i.e. the minimum wage bill a firm has to pay for the workers it has).

One issue with using the parameters from a two-way fixed effect model in our calculation of quasi-rents is that any estimation error in these estimates will be present in our measure. This will lead our estimate of the pass-through of quasi-rents to be biased towards zero.

3.2 Summary Statistics

The aggregate relationship between firm performance and wages reflects co-movements over time (wages rise when overall economic performance is good), cross-sectional covariance (high-performing firms and industries pay higher wages), and firm-level pass-through of performance improvements as higher wages. The labour income share is one summary measure of this aggregate relationship. Based on our wage and firm performance measures, we construct a measure of the labour income share and decompose the LIS into the following components:

$$\frac{wL}{VA} = \frac{bL}{VA} + \frac{(w-b)L}{VA} = \frac{bL}{VA} + \frac{(w-b)L}{QR} * \frac{QR}{VA} \tag{3}$$

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[22] The positive adjustment to the firm-earnings premiums is balanced by a reduction in the reservation wage, to remove any overall rent-sharing and to ensure the equation remains balanced. Any error in the size of the adjustment is assumed to be due to incorrectly classifying the overall level of rent sharing, and estimated $\ln \widehat{QR}$ will differ from true QR by a constant: $\ln \widehat{QR} = \ln(QR + \lambda QR) \approx \ln(QR) + \lambda.$ The bias will be absorbed by the intercept when $\ln \widehat{QR}$ is included in linear regressions.

[23] Some firms have a negative estimated premium, although these firms are generally very small (FTE<1). This negative premium could be the result of the firm-earnings premiums for these firms being poorly identified or from some form of compensating differential, where the non-pecuniary benefits of working at one of these firms is enough to compensate for the wage rate being below reservation. Given the relatively small FTE counts in these firms, we suggest the former is the most likely explanation for apparent negative premiums.

Where $\frac{wL}{VA}$ is the labour income share (total wages divided by total value added), $\frac{bL}{VA}$ is the reservation wage share of value added (total reservation wages divided by total value added), $\frac{(w-b)L}{VA}$ is total wage premiums as a share of total value added. The wage premium share of value added can be further decomposed into the share of rents captured by labour $\frac{(w-b)L}{QR}$ and the rent share of value added $\frac{QR}{VA}$. Figures 6-8 plot the time series of the above components.

Figure 6 plots our measure of the labour income share and the reservation wage share of value added. We also plot the official measure of the labour income share for comparison.[24] Our measure of the LIS sits above the official measure but follows a similar trend, giving us confidence that what is happening in our sample is representative of what is happening in the economy in general.[25]

The labour income share is increasing over the first half of the sample period, before flattening out at just over 45%. This indicates that wage growth exceed growth in value added per worker over this period. The reservation wage share follows a similar pattern to the overall LIS, increasing from 29% in 2002 to 33% in 2010 and staying around that level for the rest of the period. The difference between the LIS (dashed line) and the reservation wage share is the wage premium share, which ranges between 12% and 14% of value added.

[24] The official LIS is calculated as compensation of employees divided by the sum of compensation of employees, gross operating surplus, and gross mixed income from StatsNZ's System of National Accounts GDP(I) series. The data is for the private market sector.

[25] A key reason why the level of our measure of the LIS is above that of the official measure is the treatment of working-proprietor only firms. These firms contribute to value added but not to compensation of employees or total wages. Therefore our measure of total value added will be lower than the official measure, resulting in a higher LIS.

Figure 6: Share of wages and reservation wages in value added

Figure 7 plots the wage premiums as a share of total value added ($\frac{(w-b)L}{VA}$) and as a share of total quasi-rents ($\frac{(w-b)L}{QR}$). While both follow a similar pattern over time, rising until 2010 and then declining, the magnitude of the changes is very different. The premium share of value added is essentially flat over the sample period, with a total range of around one percentage point. The wage premium share of rents has a range of around 8 percentage points. The reason for such a difference in the variation between the two series is due to variation in the share of rents in value added. This is shown in Figure 8. The share of quasi-rents in value added declined in the first part of the period from around 43% in 2002 to 35% in 2010. This decline almost completely offsets the increase in the premium share of rents to leave the premium share of value added slightly higher in 2010 than in 2002. The share of rents in value added increases from 2010 to sit at 39% in 2018.

Figure 7: Wage premiums as a share of total value added and total rents

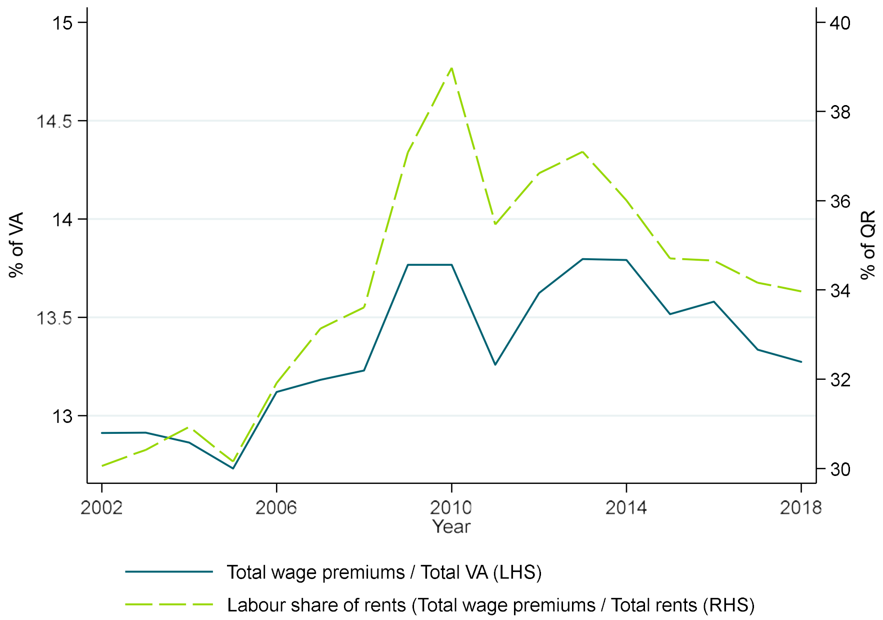


Figure 8: Share of quasi-rents in value added

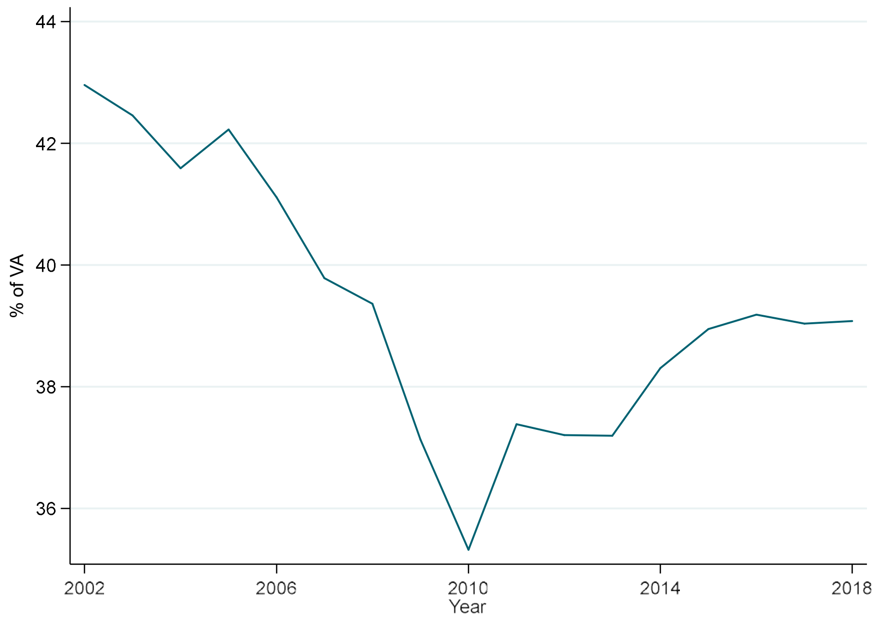


Table 2: Summary statistics – characteristics of the firm of the average worker

	Arithmetic mean	Geometric mean	Mean of log
VApw	$137,400	$105,300	11.56
	($301,600)	($102,100)	(0.68)
QRpw	$52,900	$35,400	10.47
	($246,900)	($75,800)	(1.14)
Avg. wage estimate	$64,200	$60,800	11.02
	($23,200)	($23,200)	(0.32)
FTE employment	930	85	4.44
	(2,106)	(911)	(2.46)
Adj. FTE ratio	81.9%	79.1%	-
	(17.7)	(27.7)	-
Reservation wage	$44,000	$42,800	10.66
	($10,900)	($11,400)	(0.24)
Wage premium*	$20,200	$18,000	0.35
	($15,600)	($3,400)	(0.17)
Profit pw	$34,500	$23,100	10.05
	($244,400)	($67,400)	(1.37)
Avg. tenure (months)	47.6	42.6	3.75
	(22.5)	(26.4)	(0.48)
Firm age (years)	22.8	15.6	2.75
	(20.5)	(24.3)	(0.94)
K-VA ratio	29.6%	16.9%	-
	(627.2)	(30.1)	-
M-GO ratio	43.9%	36.9%	-
	(22.3)	(38.7)	-
N	1,409,835	1,409,835	1,409,835
N Firms	266,580	266,580	266,580

Notes: Standard deviation in parentheses. The number of observations and number of firms have been randomly rounded to base 3 for confidentiality purposes. All observations are weighted by FTE employment. The geometric mean and log mean are conditional on the value of the variable being positive. Geometric means are calculated by taking the exponent of the average of the logs.

*: The reported wage premium is calculated as the difference between the average wage and the reservation wage in each column.

Table 2 presents summary statistics for our estimation sample. We present the arithmetic mean, the geometric mean, and the mean of the logs of our variables. All observations are weighted by FTE employment so the statistics describe the characteristics of the firm of the average worker. Average value added per worker (VApw) is between $105,000 and $138,000, depending on the measure used. Quasi-rents per worker (QRpw) are between $35,000 and $53,000, while the average firm wage is between $60,000 and $64,000. The distributions of the firm performance measures are much wider than that of wages across firms. The geometric standard deviation is roughly equal to the mean for value added per worker, more than twice the mean for quasi-rents, while the standard deviation of the wage distribution is less than half of the mean. The average firm reservation wage is between $42,000 and $44,000, or 68%-70% of the observed wage. Wage premiums average between $18,000 and $20,000.

The average worker works at a firm with 85 FTE employees (based on the geometric mean). There is a large difference between the arithmetic and geometric means of FTE employment, indicating the influence of some very large firms. We have sufficient wage information for around 80% of the FTE at the average firm (adj. FTE ratio).

We next look at the evolution of the distribution of wages, value added per worker, wage premiums, and quasi-rents per worker over time. Figure 9 plots the employment-weighted (between-firm) standard deviations of the (log) wages and value added per worker over the sample period, while Figure 10 plots the standard deviations of wage premiums and quasi-rents per worker.

Between-firm wage dispersion generally follows a cyclical pattern, although there is a long-run decline in the dispersion. The standard deviation reaches a minimum during the GFC, where job-losses were concentrated at the lower end of the wage distribution. Dispersion rose during the early stages of the recovery from the GFC as these workers re-entered employment, before beginning to decline during the latter part of the sample period. The total decline in wage dispersion is approximately two log-points. New Zealand is one of very few OECD countries to see a reduction in wage dispersion over this period (e.g. Barth et al. (2016) for the US, Berlingieri et al. (2017) and Criscuolo et al. (2020) for a selection of OECD countries). The dispersion in value added per worker increased markedly during the GFC due to the uneven impact of the GFC. Dispersion in value added has been trending upwards post-GFC. The correlation between wage and value added dispersion is relatively low in the employment-weighted distributions.[26]

[26] In the unweighted distribution, dispersion in both value added per worker and wages has declined.

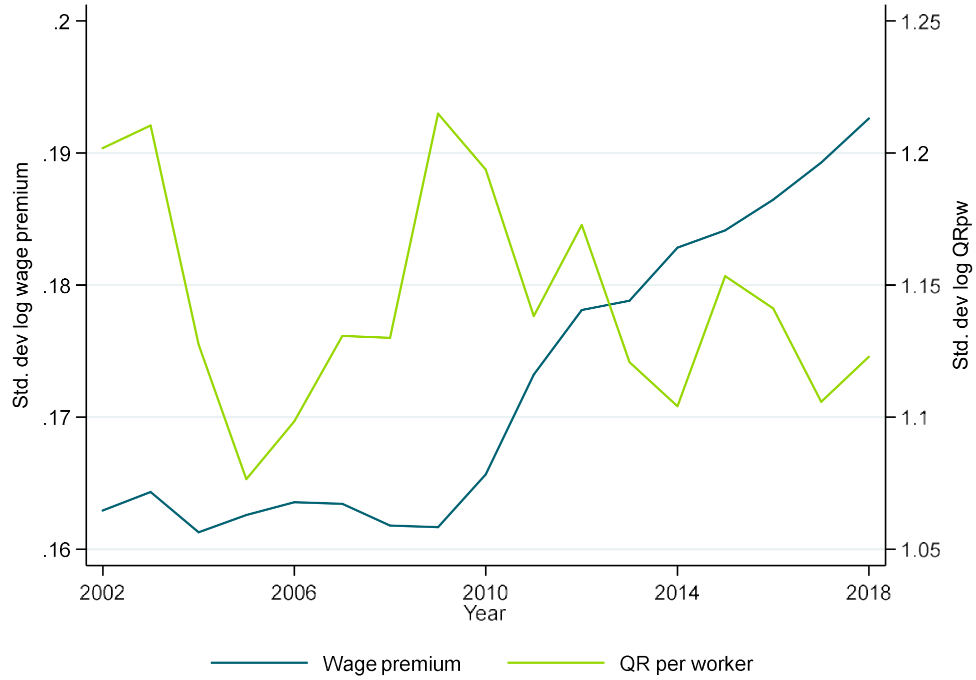
Figure 9: Standard deviation of average wages and VA per worker over time (FTE weighted)

Dispersion in wage premiums across firms is relatively stable during the early part of the sample period. From 2008, dispersion in firm premiums increases to be approximately three log-points higher in 2018. Dispersion in quasi-rents shows no obvious trend over the sample period. Dispersion declined from a peak just prior to the GFC, before beginning a gradual decline from 2010 onwards. Again, the correlation between dispersion in wage premiums and quasi-rents is low.

Figures 9 and 10 together suggest the weight of employment is more evenly spread across the distribution of value added and firm premiums. A number of studies find a decline in job-to-job transitions and job starts and ends in New Zealand (e.g. Ball et al. 2019; Maré 2018; Coleman & Zheng 2020). We also find a decline in the number of job starts and job ends.[27] Our results here suggest that high-skilled workers are not as concentrated in high rent/high premium firms to the same extent during the latter half of our sample period (2010-2018).

[27] See Figure A1 in Appendix A.

Figure 10: Standard deviation of wage premiums and quasi-rents per worker over time (FTE weighted)



4 Empirical methodology

4.1 Baseline model

We begin our examination of the extent of pass-through from rents to wages by estimating models similar to those used in the international literature. We begin with the equation:

$$\mathbf{\ln w_{jt} = \gamma \ln Rent_{jt} + X_{jt}\beta + \varepsilon_{jt}} \quad (4)$$

where $\varepsilon_{jt} = \theta_t + \psi_j + e_{jt}$. The subscript *j* indexes firms and *t* indexes years. $Rent$ is either value added per worker or quasi-rents per worker, X is a vector of control variables, and ε is a composite error term, comprising time effects (θ_t), firm unobservables (ψ_j), and a stochastic component (e_{jt}).[28] γ, the pass-through elasticity, is our coefficient of interest.

Estimating equation 4 presents a number of empirical challenges, including controlling for heterogeneity, measurement error and transitory fluctuations, and endogeneity. The aggregate relationship between firm performance and wages reflects co-movements over time (wages are higher in years where performance is good), cross sectional covariance (high-performing firms and industries pay higher wages) and firm-level pass-through of performance improvements as higher wages. It is the third of these that we aim to isolate.

Heterogeneity

We present estimates of equation 4 that include alternative sets of firm controls, workforce controls, fixed effects (detailed industry and firm), and time effects. The firm observables we include are firm age, average tenure, firm size, the fraction of FTE employment included in our wage measure, and the share of intermediates in gross output.[29] Time effects control for macro shocks that affect both performance and wages. Industry dummies control for the fact that some industries have higher performance and higher wages, possibly a result of differences in production technology or differences in internationalisation. Firm fixed effects control for firm-specific, time invariant unobservables, such as differences in production technology, management and human resource practices, or worker quality.

Firms also differ in the composition of their workforces along dimensions such as age, gender, qualifications, and skill. Our analysis in this paper is at the firm level so we cannot control for individual age, gender, or skill directly. Firm fixed effects will control for permanent differences in worker quality across firms but will not capture differential changes in composition over time. We control for differences in workforce

[28] Some firms have negative values for value added and quasi-rents per worker. We account for these observations by including a dummy variable and the term $-\ln(\text{abs}(\text{Rent}))$, rather than dropping the observation entirely.

[29] The share of intermediates in gross output is included to allow for differences in input substitutability.

composition by including the sum of the firm-average covariate index and firm-average worker fixed effects from the 2-way fixed effect model ($\bar{a}_j + \bar{Z}_{jt}$ using the notation of equation 1, where bars denote firm averages). The covariate index controls for average age and gender composition and the average worker fixed effect is a proxy for worker skill. Using this as our workforce control is equivalent to running a regression using our estimate of the wage premium as the LHS where the coefficient on the workforce composition variable is equal to one.

Measurement error and transitory shocks
Firm performance measures are known to be highly variable. In our sample the standard deviation of log value added is more than twice that of average wages, while the standard deviation of log quasi-rents is more than three times that of wages. This means that current performance may be only weakly related to underlying performance. This could be due either to measurement error or temporary year-to-year volatility. The presence of either transitory shocks or measurement error induces a correlation between our performance measures and the error term and will result in our estimate of γ being biased towards zero. A common solution to this is to use instrumental variables to remove the influence of measurement error and short-term transitory fluctuations. We follow the literature and use instruments to address the attenuation bias caused by measurement error and/or transitory fluctuations.

We use the ratio of industry-level output and input price indices and the input price index interacted with the current firm-level ratio of intermediates to capital as instruments. The ratio of output and input price indices tells us about the evolution of mark-ups across industries over time, while the ratio of intermediates to capital tells us about the importance of intermediates in the firm input mix. These instruments isolate price-driven movements in rents, which are more likely to represent permanent changes that are potentially more likely to be passed on to workers.

As a robustness check, we test two further specifications as alternative methods for controlling for firm heterogeneity and transitory fluctuations. The first is a differenced specification with increasingly longer differences. This eliminates any permanent, firm specific differences, as in a firm fixed effect specification. Juhn et al. (2018) show that increasing the length of the difference places relatively more weight on changes in the permanent component of performance, so the estimated coefficient should increase with the length of the difference. The second is a correlated random effects specification where we include the firm-level average of the RHS variables as well as the current-year deviation from the average. This specification allows us to examine the pass-through of both the permanent and transitory components of firm performance in a single model.

Studies that employ instrumental variables typically find that pass-through is stronger in their IV models than in their OLS models (Bell et al. 2019; Van Reenen 1996; Card et al. 2014, 2018; Kline et al. 2019; Abowd & Lemieux 1993; Juhn et al. 2018). This suggests that the instruments used are removing the influence of measurement error or transitory fluctuations and better isolating the permanent component of firm performance. The instruments used in the literature include measures of innovation

Van Reenen 1996; Kline et al. 2019; Hildreth 1998), export and import prices (Abowd & Lemieux 1993; Martins 2009), or measures of physical productivity (Carlsson et al. 2016). Papers that estimate dynamic panel models use lags as instruments (e.g Bell et al. 2019; Hildreth & Oswald 1997). Others use firm performance in the same industry but other regions as instruments (e.g. Card et al. 2014; Barth et al. 2016). All of these instruments are intended to remove the influence of transitory shocks to firm performance, such as temporary or localised demand shocks.

Simultaneity

There are also a number of simultaneity concerns. The first concerns the nature of shocks to firm performance. These shocks often have a market or industry component to them, which will tend to raise the market wage rate. This generates a positive correlation between firm performance and wages even in the absence of any rent sharing, leading to the overestimation of the extent of pass-through. Time effects alleviate the impact of general macroeconomic shocks, but do not fully account for industry-specific shocks. The price-based instruments (e.g. export and import prices, exchange rates) tend to be measured at a more aggregate level so these may fail to correct for this positive bias. However, it's unclear how important this source of bias is in comparison with the other sources.

The second can occur when bargaining is inefficient. If bargaining is efficient, firms and workers are bargaining over the sharing of any surplus (i.e. wages and profits). Employment (and hence output) has already been decided through the standard profit maximisation problem of the firm. If bargaining is inefficient, then higher wages will induce firms to reduce employment, thereby reducing output and rents. Inefficient bargaining will generate a negative correlation between wages and measures of performance, leading to an underestimate of the extent of pass-through. It's not obvious that there is a relationship between prices and the efficiency of bargaining, so our price-based instruments should reduce this issue if present.

Other endogeneity issues

A further potential endogeneity issue is the presence of efficiency wages (Akerlof & Yellen 1986). If firms are paying efficiency wages, we would expect wages to be driving firm performance, rather than firm performance driving wages. This will generate a positive correlation between performance and wages and will lead us to overstate the extent of pass-through. Using quasi-rents, which accounts for differences in worker quality (if worker quality is driving performance) and directly controlling for workforce composition reduce the impact of this potential source of bias.

Another potential source of bias is the presence of firm-specific amenities that affect the willingness of a worker to work at the firm (e.g. Kline et al. 2019). A positive shock to these amenities enables firms to pay lower wages, which will allow it to expand employment. This will tend to lower the average product of labour and thereby induce a positive correlation between wages and rents, leading to an overestimate of the pass-through elasticity.

An improvement in amenities could also induce a negative correlation between rents and wages. The literature on management practices shows that good management practices raise worker morale and productivity (e.g. Bloom & Van Reenen 2007). The boost to worker morale allows firms to lower wages while also boosting rents. This leads to a negative correlation between wages and rents resulting in an underestimate of the extent of pass-through. Our price-based instruments should reduce the influence of shocks to firm amenities as it is unlikely that these shocks are related to prices.

4.2 Worker sorting vs. rent-sharing

Our workforce composition variable has the effect of removing the influence of worker sorting on our estimates of the pass-through elasticity. However, we are interested in the extent to which worker sorting influences pass-through and how worker sorting has changed over time. To get clearer insights on these questions, we drop the workforce composition variable and implement the decomposition used in Card et al. (2018). This technique makes use of the fact that, if the 2-way fixed effect specification is an appropriate specification for wages, then equation 4 can be written as:

$$\bar{a}_j + \phi_j + \bar{Z}_{jt}\lambda + \tau_t + \bar{\xi}_{jt} = \gamma \ln Rent_{jt} + X_{jt}\beta + \varepsilon_{jt} \tag{5}$$

Where bars denote firm averages. To decompose γ, we estimate the two equations:

$$\bar{a}_j + \bar{Z}_{jt}\lambda + \tau_t = \gamma_R \ln Rent_{jt} + X_{jt}\beta + \varepsilon_{jt} \tag{6}$$

$$\phi_j + \xi_{jt} = \gamma_P \ln Rent_{jt} + X_{jt}\beta + \varepsilon_{jt}$$

where $\gamma = \gamma_R + \gamma_P$.

$\bar{a}_j + \bar{Z}_{jt}\lambda + \tau_t$ is our estimate of the (log) average reservation wage and $\phi_j + \xi_{jt}$ is our estimate of the firm wage premium. The coefficient γ_R then measures the correlation between rents and reservation wages, capturing the sorting of high-skilled workers (with higher reservation wages) to high-rent firms (if $\gamma_R > 0$). γ_P measures the association between the firm-specific component of wages and firm rents and provides a cleaner estimate of the rent-sharing elasticity.

Controlling for workforce composition and the Card et al. (2018) decomposition are equivalent specifications when the coefficient in equation 4 on the worker fixed effects plus covariate index is equal to one. In some specifications, we regress the wage premium on the full set of firm observables and a firm fixed-effect, including our measure of workforce composition. This relaxes the restriction that the coefficient on the workforce composition variable is equal to one when we use the observed wage as the LHS variable and captures the correlation between worker quality and firm pay premiums.

Imperfect normalisation of the reservation wage, as described in section 3.1.3 may result in biased estimates of γ_R and γ_P due to a correlation of the estimated reservation wage (and consequently the estimated premium) with log rents. Similarly, because the estimated reservation wage is used in the construction of the quasi-rent variable, any estimation errors on the reservation wage will be negatively correlated with estimated quasi-rents, and will also result in a negative correlation between the estimated premium and estimated quasi-rents. The use of instrumental variables to instrument for the log of quasi-rents will remove any of these biases in the estimation of γ_R and γ_P

4.3 Time varying pass-through

We estimate a simple extension to equation 4 to test for changes over time in the extent of pass-through and the contributions of worker sorting and rent sharing. Specifically, we estimate:

$$\ln w_{jt} = \gamma_t \ln Rent_{jt} + X_{jt}\beta + \varepsilon_{jt} \tag{7}$$

γ_t is a vector of coefficients on our rent measures interacted with year dummies. We do this for our main specifications and for the Card et al. (2018) decomposition. This allows us to test whether worker sorting or rent-sharing have become more or less important in explaining pass-through over time. We estimate time-varying pass-through estimates in both our OLS model and our IV model. In our IV model, we interact our instruments with year dummies.

5 Results

We now present our estimates of the extent of pass-through from performance to wages, based on equation 4. We begin by examining overall pass-through and the extent to which this reflects cross-sectional differences across firms and within-firm pass-through of firm performance. We then turn to our decomposition results to examine the role of worker sorting and rent sharing in explaining pass-through. Finally, we look at whether pass-through, worker sorting, and rent-sharing have changed over time.

5.1 Baseline results

We begin by presenting OLS estimates of equation 4, adding in sets of control variables sequentially to examine the impact these have on the estimated pass-through elasticity. These are presented in Table 3. Column 1 presents the bivariate pass-through elasticity and contains only year dummies as controls. Column 2 introduces our control for workforce composition, the sum of the average worker fixed effect and average covariate index from a two-way fixed effect model. Column 3 augments Column 1 by including the set of firm observables: firm FTE employment, the fraction of FTE included in our wage measure, average tenure, firm age, the intermediate share of gross output, and ANZSIC 3-digit industry dummies. Column 4 includes a firm fixed effect, and column 5 includes the full set of firm and worker controls.

There is relatively strong cross-sectional covariance between wages and firm performance. The coefficients in the first column of Table 3 imply that a 10% higher level of VAPW is associated with 2.5% higher wages, and a 10% higher level of QRPW is associated with 1.4% higher wages. About half to two thirds of this effect is due to higher-performing firms having more skilled workers (column 2) or having other observable characteristics that are correlated with both higher wages and better performance (column 3). Adding firm fixed effects in column 4 further reduces the estimated effect, reflecting the fact that the within-firm covariance of firm performance and wages is relatively weak. Most of the effect in column 3 is due to better performing firms paying higher wages than observationally similar firms, rather than because firms pay higher wages when performance improves. The final column of Table 3 presents estimates of the within-firm covariance, controlling also for worker composition in the firm. This least restrictive specification shows that a 10 % higher VAPW is associated with wages that are only 0.32% higher, and a 10% higher QRPW is associated with wages that are only 0.13% higher.

Table 3: Wage pass-through: Value added and Quasi-rents (OLS Estimates)

Dependent Variable: ln(wage)	Year effects only	Worker controls	Firm observables	All Firm controls	Worker & firm controls
	(1)	(2)	(3)	(4)	(5)
			VApw		
Ln(VApw)	0.254***	0.0887***	0.130***	0.0439***	0.0318***
	(0.00913)	(0.00765)	(0.00499)	(0.00250)	(0.00173)
R-squared	0.462	0.805	0.784	0.957	0.970
Within Adj R2	0.309	0.749	0.382	0.0453	0.325
Fixed effects	Year	Year	Year, industry	Year, firm	Year, firm
			QRpw		
Ln(QRpw)	0.137***	0.0495***	0.0547***	0.0126***	0.0126***
	(0.00604)	(0.00500)	(0.00202)	(0.000999)	(0.000918)
R-squared	0.415	0.805	0.770	0.956	0.970
Within Adj R2	0.248	0.749	0.341	0.0340	0.326
Fixed effects	Year	Year	Year, industry	Year, firm	Year, firm
Observations	1,409,835	1,409,835	1,409,835	1,351,134	1,351,134
N firms	266,580	266,580	266,580	207,879	207,879

Notes: Standard errors, in parentheses, are clustered at the firm level. ***, **, and * denote statistical significance at the 1%, 5%, and 10% level, respectively. All models include time dummies. Firm observable control variables are firm size (FTE), average tenure of workers, firm age, intermediate input share of gross output, and ANZSIC level 3 industry dummies. Worker controls are the firm average worker fixed effect plus the firm average covariate index from a 2-way fixed effect model. The number of observations and the number of firms have been randomly rounded to base 3 for confidentiality reasons. All equations include a term for firms with negative value added or quasi-rents. Regressions performed using the reghdfe command in Stata 16 (Correia 2017). Columns 4 and 5 include firm fixed effects. Firms with only one annual observation are dropped from these specifications.

Appendix B presents some robustness tests for the results in columns 4 and 5 of Table 3 using some alternative specifications. Table B1 presents results from a differenced model and Table B2 presents results from a correlated random effects model. In the differenced model, we see that the estimated pass-through elasticity increases with the length of the difference. This is consistent with the results of Juhn et al (2018), who show that increasing the length of the difference increases the relative weight on the permanent component of shocks to firm performance. In the correlated random effects model, we see that within-firm pass-through is weaker than across-firm pass-through. Both tables show that changes in workforce composition are more important in explaining pass-through of value added than in explaining pass-through of quasi-rents. Overall, these results are consistent with those in Table 3.

As noted in section 4.1 above, OLS estimates of equation 4 as presented in Table 3 may be biased due to a range of mis-specifications, including endogeneity of firm performance, and attenuation bias from measurement error or transitory performance volatility. To test the sensitivity of pass-through estimates to these issues, we re-estimate the final three columns of Table 3 using an IV estimator. The instrumental variables that we use are described in section 4.1 and are constructed from industry level input and output prices interacted with firm-level capital-intermediates input mix. These instruments are correlated with firm performance but unlikely to be correlated with idiosyncratic wage variation. The IV estimates are presented in Table 4.

Table 4: Wage pass-through: Value added and Quasi-rents (IV Estimates)

Dependent Variable: ln(wage)	Firm observables (T3, col 3) (3’)	All firm controls (T3, col 4) (4’)	Worker & firm controls (T3, col 5) (5’)
		VApw	
Ln(VApw)	0.187***	0.134***	0.116***
	(0.0475)	(0.0461)	(0.0291)
R-squared	0.368	-0.041	0.244
Overid: p-value (H0: overidentified)	0.352	0.250	0.120
UnderID: KP p-value (H0: not identified)	0.000	0.000	0.000
Weak IV (H0: instruments are weak)	4.123	11.994	11.666
Observations	1,384,770	1,326,636	1,326,636
N firms	262,050	203,919	203,919
		QRPW spec	
Ln(QRpw)	0.111***	0.0909**	0.0684***
	(0.0214)	(0.0391)	(0.0256)
R-squared	0.285	-0.376	0.128
Overid: p value (H0: identified)	0.499	0.180	0.157
UnderID: KP p value (H0: not identified)	0.000	0.000	0.000
Weak IV (H0: instruments are weak)	2.283	3.423	3.432
Observations	1,130,397	1,072,656	1,072,656
N firms	228,495	170,754	170,754

Notes: See notes to Table 3. Overid is the p-value for the Hansen’s J test of over-identifying restrictions. UnderID is the p-value of the Kleibergen-Paap rk test for under-identification (Kleibergen & Paap 2006). Weak IV is the Kleibergen-Paap Wald rk F statistic. Firms with negative rents or value added are dropped from this estimation. Estimation carried out using 2SLS using the ivreghdfe command in Stata 16 (Correia, 2017; Baum et al. 2010)

Consistent with most international studies of pass-through, the IV estimates of the pass-through coefficients are considerably larger than those based on OLS. This is likely to be due to the relatively strong influence of transitory variation and measurement error in the firm performance measures, which biases the OLS estimates towards zero. In the most restrictive specification, the pass-through coefficient increases roughly four-fold, implying that a 10 percent increase in VApw is associated with 1.3 percent higher wages, and a 10 percent increase in QRpw is associated with a 0.6 percent higher wage.[30]

Table 5: Interpreting the size of estimates

	Firm observables (T3, col 3) (1)	All firm controls (T3, col 4) (2)	Worker & firm controls (T3, col 5) (3)
OLS ESTIMATES			
		VApw	
Coefficient	0.130	0.044	0.032
Implied impact of a $10,000 increase in VApw	$590	$200	$140
		QRpw	
Coefficient	0.055	0.013	0.013
Implied impact of a $10,000 increase in QRpw	$620	$140	$140
IV ESTIMATES			
		VApw	
Coefficient	0.187	0.134	0.116
Implied impact of a $10,000 increase in VApw	$860	$614	$532
		QRpw	
Coefficient	0.111	0.0909	0.068
Implied impact of a $10,000 increase in QRpw	$1,270	$1,040	$783

In order to aid the interpretation of the coefficients, Table 5 shows what the coefficients imply about the dollar increase in annual wages in response to a $10,000 increase in either value added per worker or quasi-rents per worker. A $10,000 increase

[30] The instruments are rather weak (Kleibergen-Paap Wald rk F values between 4 and 12 for the VApw spec and between 2 and 3 for the QRpw spec), but satisfy an over-identification test (p value of 0.2 to 0.9) indicating that they are not correlated with idiosyncratic wage variation, and also pass an under-identification test (p value of 0) indicating that they are independently correlated with the measures of firm performance.

amounts to about 8 percent of mean VAPW, or about 20 percent of mean quasi-rent per worker, or 5 and 4 percent of one standard deviation respectively. While the estimated elasticities on quasi-rents are lower than those on value added, they imply stronger dollar pass-through. The first column shows an annual wage increase of around $600 for value added and $650 for quasi-rents, or about 6c to 6.5c per dollar of rents, based on our OLS estimates. IV estimates suggest an annual wage increase of between $800 for value added and $1,300 for quasi-rents, or 8c to 13c per dollar of rents. The IV estimates of our least restrictive specification show that workers receive 5c in the dollar of improvements in value added, and 8c in the dollar of improvements in quasi-rents. This is consistent with quasi-rents being a better measure of the resources available to be split between wages and profits. A dollar increase in quasi-rents is a larger proportional change than a dollar increase in in value added, and with the same wage variation, necessarily results in a lower estimated elasticity on quasi-rents.

5.2 Worker sorting vs. rent sharing

Our preferred pass-through estimates are those in column 5 and 5' in Tables 3 and 4. These control for changing worker composition, removing the influence of worker sorting on the estimated coefficients.

However, separating the contributions of sorting and sharing requires us to drop the worker composition controls.[31] The focal specification for our analysis of sorting and sharing is therefore the specification in columns 4 and 4' in tables 3 and 4. We also report estimates based on specifications that omit firm fixed effects (columns 3 and 3'). Based on these specifications, we implement the Card et al. (2018) decomposition of the pass-through parameter in to worker sorting and rent-sharing contributions, as outlined in section 4.2 above.

Table 6 presents the decomposition results of our OLS estimates from Table 3. Columns 1 and 2 decompose the estimate from the firm observables specification (column 3 of Table 3), columns 3 and 4 decompose the estimate from the firm fixed effect specification (column 4 of Table 3) This decomposition replaces the observed wage as the LHS variable and replaces it with either our estimate of the reservation wage, to examine sorting, or the firm wage premium, to examine rent-sharing. The specification in column 5 replaces the observed wage with the wage premium as the left-hand side variable and includes the full set of controls, including workforce composition.

Worker sorting explains approximately 50% of our estimated value-added pass-through elasticity in our firm observables specification. This is similar to the results in Card et al. (2018). Most of this sorting is between firms. In our firm fixed effect specification, sorting accounts for about one third of the overall pass-through estimate. Sorting

[31] The worker composition controls are collinear with the reservation wage estimate shown in equation 6, so the sorting equation is uninformative. Estimating the sharing (wage premium) equation based on column 5 or 5' yields estimates that are essentially identical to those based on column 5 or 5'.

accounts for a lower proportion of quasi-rents pass-through, which is expected as quasi-rents are net of the reservation wage bill so worker quality is partly accounted for. Sorting accounts for 37% of pass-through in the firm observables specification and essentially none of the pass-through in the firm fixed effects specification. It appears that all sorting on quasi-rents is cross-sectional – good workers don't tend to move to firms with higher quasi-rents, they already tend to be working in these firms.

Table 6: Sorting or sharing (OLS Estimates)

	Firm observables (T3, col 3)		All firm controls (T3, col 4)	
	(1)	(2)	(3)	(4)
VARIABLES	Ln(Reservation wage)	Ln(Wage premium)	Ln(Reservation wage)	Ln(Wage premium)
		VApw		
Ln(VApw)	0.0645***	0.0659***	0.0149***	0.0290***
	(0.00289)	(0.00250)	(0.00165)	(0.00166)
R-squared	0.823	0.550	0.967	0.873
Within Adj R2	0.306	0.263	0.143	0.038
		QRpw		
Ln(QRpw)	0.0205***	0.0341***	2.46e-06	0.0126***
	(0.00106)	(0.00121)	(0.000526)	(0.000926)
R-squared	0.813	0.548	0.967	0.874
Within Adj R2	0.269	0.259	0.138	0.042
Observations	1,409,835	1,409,835	1,351,131	1,351,131
N Firms	266580	266580	207,876	207,876

Notes: See notes to Table 3

Table 7 presents the decomposition of our IV estimates. The point estimate for worker sorting on value added in the firm observables specification is similar to that in Table 6, though imprecisely estimated. All of the difference between the OLS and IV estimates is attributed to sharing. Adding firm fixed effects reduces the coefficient in the sorting equation by two thirds, while the coefficient in the sharing equation is marginally smaller. The results indicate that 10% higher value added per worker is associated with a 1.1% increase in wages, which is similar to the full model that controls for workforce composition.

Worker sorting is more important in explaining pass-through of quasi-rents in our IV estimates than in our OLS estimates. Sorting accounts for approximately 45% of the estimated quasi-rent pass-through in the firm observables specification, compared with 37% in the OLS equivalent. Again, much of this sorting appears to be cross-sectional, with the importance of sorting declining when firm fixed effects are included. Our estimates for worker sorting are larger than those in the OLS estimates, suggesting that our instruments are removing the influence of the estimation error in the reservation wage bill in our quasi-rents variable.

Table 7: Sorting or sharing: (IV Estimates)

	Firm observables (T4, col 3')		All firm controls (T4, col 4')	
	(1)	(2)	(3)	(4)
VARIABLES	Ln(Reservation wage)	Ln(Wage premium)	Ln(Reservatio n wage)	Ln(Wage premium)
		VApw		
Ln(VApw)	0.0601	0.127***	0.0224	0.111***
	(0.0427)	(0.0367)	(0.0465)	(0.0307)
R-squared	0.308	0.215	0.142	-0.065
Overid: p-value (H0: over-identified)	0.248	0.795	0.769	0.141
UnderID: KP p-value (H0: not identified)	0.000	0.000	0.000	0.000
Weak IV (H0: instruments are weak)	4.123	4.123	11.994	11.994
Observations	1,384,770	1,384,770	1,326,636	1,326,636
N Firms	262,050	262,050	203,919	203,919
		QRpw		
Ln(QRpw)	0.0512***	0.0596***	0.0269	0.0640**
	(0.0150)	(0.0120)	(0.0294)	(0.0250)
R-squared	0.239	0.227	0.028	-0.190
Overid: p-value (H0: over-identified)	0.210	0.877	0.596	0.201
UnderID: KP p-value (H0: not identified)	0.000	0.000	0.000	0.000
Weak IV (H0: instruments are weak)	2.283	2.283	3.423	3.423
Observations	1,130,397	1,130,397	1,072,656	1,072,656
N Firms	228,495	228,495	170,754	170,754

Notes: See notes to Table 4

5.3 Time variation in pass-through, worker sorting, and rent sharing

We now turn to the question of whether the extent of pass-through changes over time. We estimate a pass-through elasticity for each year from 2002-2018 and also produce the Card et al. (2018) decomposition of these pass-through estimates to examine changes in the importance of worker sorting and rent sharing.

Figure 11 plots the time variation in the pass-through of quasi-rents, based on our firm observables specification (panel A) and our firm fixed effect specification (panel B). These are based on our OLS estimates.

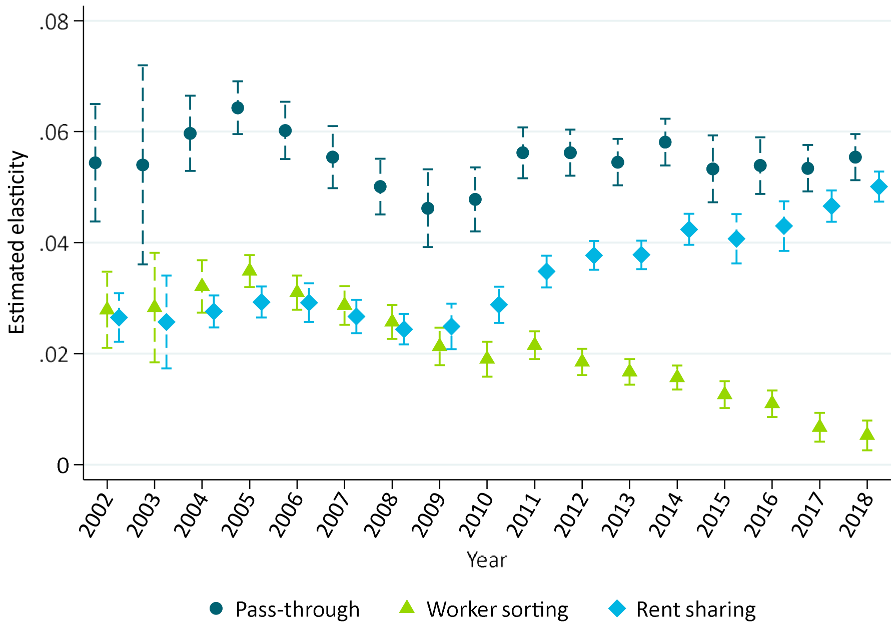
In both specifications we see a cyclical pattern in overall pass-through. We see higher pass-through elasticities in the pre-GFC period, a decline from 2008-2010 during the GFC, before they recover from 2011 onwards. In the firm fixed effect specification, pass-through returns to the pre-GFC level. In the firm observables specification, the pass-through estimates are slightly lower post-GFC than pre-GFC, although the differences are not statistically significant.

Time variation in sorting and sharing follow very different patterns to overall pass-through. The importance of rent-sharing in explaining pass-through increases substantially post-GFC, with the estimated coefficient doubling in the firm observables specification (to explain over 90% of pass-through) and increasing nearly 4-fold in the firm fixed effect specification (to explain nearly 200% of estimated pass-through). The role of worker sorting declines dramatically, becoming negative during the latter half of the sample period in the fixed effect specification. This indicates that workers of different quality are more evenly distributed across firms with different levels of rents.

Figure B1 in Appendix B shows the time variation in the pass-through of quasi-rents, based on our IV estimates. As with Figure 11, these are based on our firm observables specification and our firm fixed effects specification and show the time variation in overall pass-through, rent sharing, and worker sorting. The patterns in the point estimates are similar to those in Figure 11, namely relative stability of overall pass-through and the declining importance of worker sorting in explaining overall pass-through. However, the confidence intervals on the estimates are very wide, making it difficult to draw strong conclusions about the evolution of pass-through, rent sharing and worker sorting from these results.

Figure 11: Time variation in pass-through of Quasi-rents (OLS estimates

Panel A: Cross-firm model (T3 col 3)



Panel B: Within-firm model (T3 col 4)

6 Conclusions and next steps

We study the extent to which firm financial performance is reflected in the wages they pay their workers in New Zealand. Our aim is to understand the patterns in pass-through both between and within firms, as well as how pass-through has changed over the period 2002-2018. We also explore the importance of worker sorting and rent-sharing in explaining the pass-through we observe. We consider two measures of firm performance – value added per worker and quasi-rents per worker. We demonstrate a method to calculate quasi-rents using the estimated components of a two-way fixed effect model for wages. Quasi-rents are measured net of capital costs and reservation wages and better approximate the resources available to be split between workers in the form of higher wages or firms in the form of profits.

We find that the pass-through elasticity for value added is higher than that for quasi-rents. Cross-sectional estimates of the pass-through elasticity are 0.13 for value added and 0.05 for quasi-rents. Permanent differences in performance between firms explain a significant part of these estimates, which partly reflects differences in workforce composition across firms. Estimates from our preferred restrictive specification, which controls for unobserved firm heterogeneity and workforce composition, are 0.03 for value added and 0.01 for quasi-rents.

Our instrumental variables estimates, which reduce the influence of measurement error and transitory fluctuations in performance, are larger than our OLS estimates, consistent with international literature. Cross-sectional estimates are 0.19 for value added and 0.11 for quasi-rents. Within-firm estimates are 0.11 for value added and 0.07 for quasi-rents.

While the estimated elasticities are larger for value added, the implied per dollar pass-through is stronger for quasi-rents. Using our instrumental variables estimates for our most restrictive specification, a dollar increase in quasi-rents is associated with an 8c increase in wages, while a dollar increase in value added is associated with a 5c increase in wages. This is consistent with our measure of quasi-rents being a better proxy for the resources available to be split between workers and business owners.

We find that worker sorting explains 35%-50% of the observed cross-sectional pass-through, depending on whether value added or quasi-rents are used as the measure of performance and the estimation method. The remainder of the pass-through is explained by rent-sharing. Much of this worker sorting is cross-sectional – we find that worker sorting explains at most 25% of the observed within-firm pass-through. Workers do not appear to be systematically moving to firms with better performance.

Estimates of pass-through are relatively stable over time, displaying a slight pro-cyclical pattern. This pro-cyclical pattern likely reflects changes in uncertainty and the volatility of firm performance over the business cycle. We might expect to see more insurance-type behaviour when uncertainty and volatility are higher, leading to lower pass-through estimates. We do see large changes over time in the relative importance of worker sorting and rent-sharing in explaining overall pass-through. The importance of worker sorting declined dramatically over the period, from explaining 50%-60% of observed pass-through to having almost no role in pass-through by 2018. In the firm fixed effects specification, the contribution of worker sorting became negative, indicating that new workers to the firms with better performance reduced the average worker quality at the firm.

Our work in this paper raises a number of questions for future research. In a forthcoming paper we will explore heterogeneity in pass-through across both firm and worker characteristics. We will also explore why any differences may exist, such as the use of performance pay, collective bargaining arrangements, the prevalence of temporary migrant labour, or interactions with other labour market or income support institutions. Future work could also examine the importance of product market versus labour market rents and whether pass-through differs where labour market rents are the more important source of rents.

Another useful line of inquiry is to explore why the importance of worker sorting has changed over time. There may have been a shift in the composition of workers who move jobs or that firm are willing to raise wages in order to reduce the likelihood that their best workers leave. There could also have been a shift in the types of firms that are creating new jobs. We document an increase in the variance of the employment-weighted distribution firm pay premiums over time, indicating that the mass of employment is shifting across the pay premium distribution.

We conclude that firm performance does matter for wages, to a limited extent. Much of the relationship is cross sectional - better performing firms pay higher wages. We also show that improvements in performance are related to increases in wages, albeit to a lesser extent. This could be due to the temporary nature of changes in firm performance, which firms may insulate their workers from. Where you work matters, but the general economic conditions prevailing at the time are also crucial in explaining wage growth.

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Appendix A – job-month coverage

Figure A1: Fraction of job-months excluded from analysis

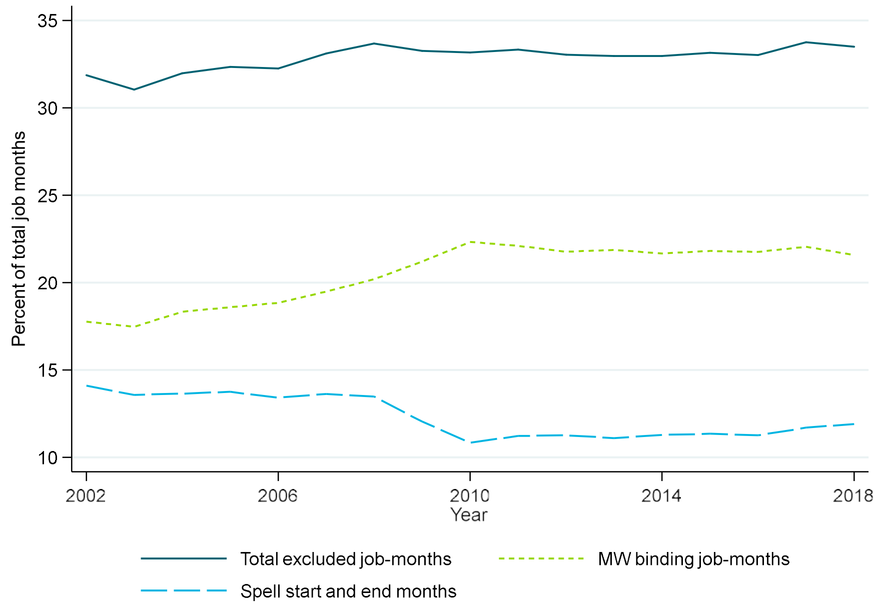


Figure A2: Percentage of total wages and employment included in analysis

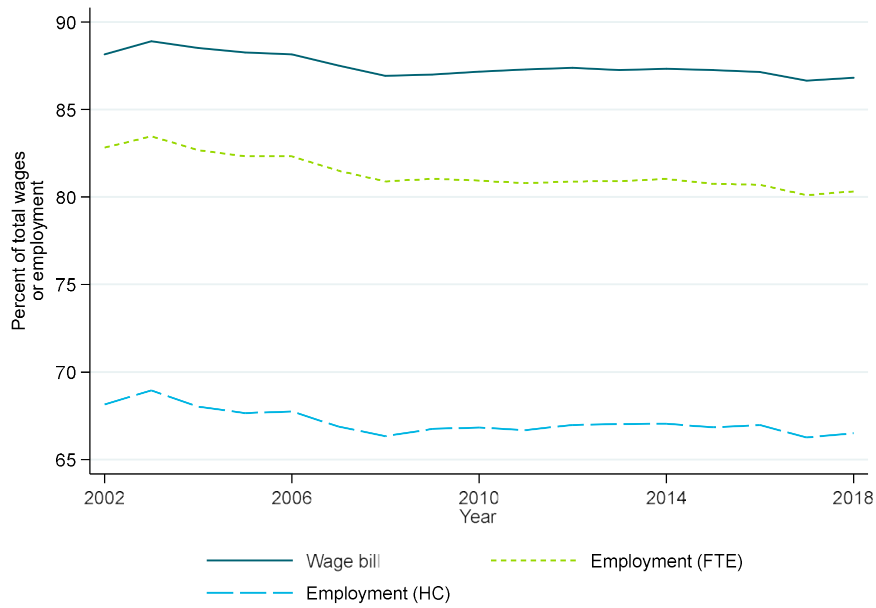
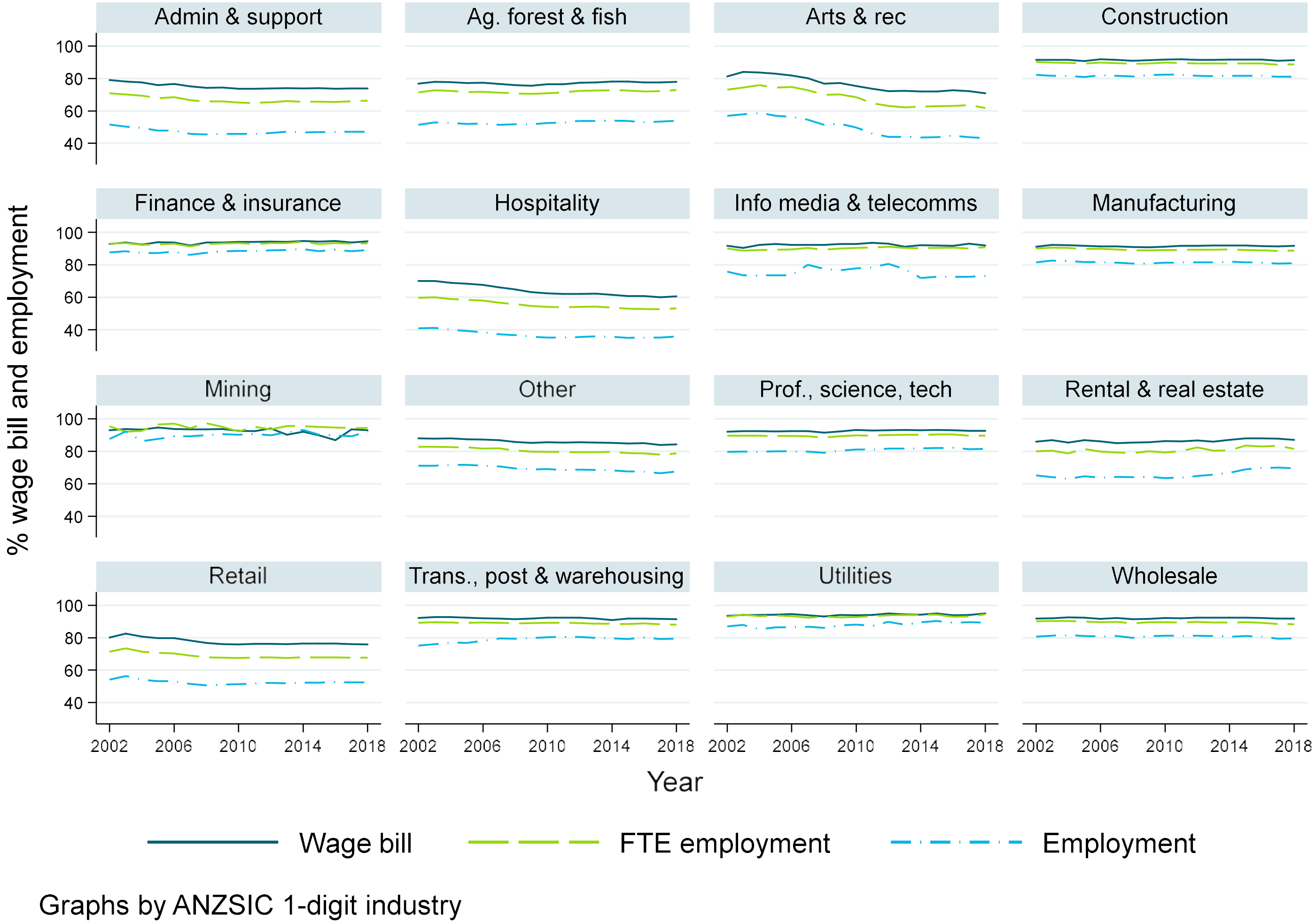


Figure A3: Percent of wages, FTE and headcount employment included by industry



Appendix B – robustness specifications

Table B1 shows the results from a differenced model using one year, three year, and five year differences. Table B2 shows the results from a correlated random effects specification. These are alternative specifications to our all firm controls (Table 3 column 4) and worker and firm controls (Table 3 column 5) specifications.

Table B1 presents pass-through estimates from a change model with a 1, 3, and 5 year difference. Columns 1-3 report the equivalent results for the all firm controls specification (Table 3 column 4) and columns 4-6 report the equivalent results for the worker and firm controls specification (Table 3 column 5). We see that the estimate increases with the length of the difference, with the estimates in the three year difference being approximately equal to the estimates in columns 4 and 5 of Table 3. This is consistent with the results of Juhn et al. (2018), who show that increasing the length of the difference increases the relative weight on the permanent component of shocks to firm performance. Estimates using a one-year difference are 50%-65% of the corresponding estimates in Table 3.

Table B2 presents the estimates from a correlated random effects model. Columns 1 and 2 show results for the equivalent of the all firm controls specification and columns 3 and 4 show the results for the equivalent of the worker and firm controls specification. We show results that include only the current-year deviation from the mean and including both the current-year deviation and the lagged deviation. The coefficient on the firm-level mean is larger than that on the deviation, confirming that cross-sectional differences in performance explain a significant proportion of observed pass-through. The estimates on the current-year deviation, the analogous estimates to those in Table 3, are generally larger than the estimates in Table 3. Part of this reflects autocorrelation in the deviation from mean. Coefficients on the lagged deviation are statistically significant and the inclusion of the lag reduces the estimate on the current-year deviation. Overall, the results in Tables B1 and B2 are consistent with those in Table 3. Within-firm pass-through is weaker than across-firm pass-through and changes in workforce composition are more important in explaining pass-through of value added than in explaining pass-through of quasi-rents.

Table B1: Wage pass-through: Value added and Quasi-rents - change model (OLS estimates)

	All firm controls (Table 3 column 4)			Worker and firm controls (Table 3 column 5)		
	(1)	(2)	(3)	(4)	(5)	(6)
	One-year change	Three-year change	Five-year change	One-year change	Three-year change	Five-year change
		VApw				
Change in ln(VApw)	0.0229***	0.0410***	0.0501***	0.0203***	0.0316***	0.0339**
	(0.00123)	(0.00242)	(0.00432)	(0.00102)	(0.00149)	(0.00223
R-squared	0.035	0.108	0.176	0.190	0.364	0.467
Within R-squared	0.016	0.049	0.078	0.174	0.322	0.404
		QRpw				
Change in ln(QRpw)	0.00713***	0.0119***	0.0147***	0.00830***	0.0130***	0.0149**
	(0.000565)	(0.00114)	(0.00181)	(0.000526)	(0.00105)	(0.00125
R-squared	0.033	0.099	0.163	0.191	0.365	0.468
Within R-squared	0.014	0.039	0.062	0.175	0.322	0.404
Observations	1,037,739	721,566	516,156	1,037,739	721,566	516,15
N Firms	199,554	138,540	102,864	199,554	138,540	102,86

Table B2: Wage pass-through: Value added and Quasi-rents - correlated random effects model (OLS estimates)

	All firm controls (Table 1 C4)		Worker and firm controls (Table 1 C5)	
Dependent Variable: ln(Wage)	(1)	(2)	(3)	(4)
		VApw		
Mean ln(VApw)	0.150***	0.150***	0.0710***	0.0712***
	[0.00845]	[0.00843]	[0.00419]	[0.00418]
$Ln(VApw)_t$ – Mean ln(VApw)	0.0531***	0.0460***	0.0411***	0.0358***
	[0.00409]	[0.00374]	[0.00273]	[0.00235]
$Ln(VApw)_{t-1}$ – Mean ln(VApw)		0.0199***		0.0150***
		[0.00257]		[0.00190]
R-squared	0.797	0.797	0.895	0.895
Within R2	0.414	0.415	0.697	0.698
	QRpw			
Mean ln(QRpw)	0.0691***	0.0691***	0.0410***	0.0410***
	[0.00215]	[0.00214]	[0.00120]	[0.00120]
$Ln(QRpw)_t$ – Mean ln(QRpw)	0.0142***	0.0129***	0.0149***	0.0136***
	[0.00135]	[0.00124]	[0.00128]	[0.00117]
$Ln(QRpw)_{t-1}$ – Mean ln(QRpw)		0.00757***		0.00738***
		[0.000838]		[0.000781]
R-squared	0.784	0.784	0.896	0.896
Within R2	0.379	0.379	0.700	0.699
Observations	1,037,739	1,037,739	1,037,739	1,037,739
N Firms	199,554	199,554	199,554	199,554

Figure B1: Time variation in the pass-through of Quasi-rents (IV estimates)

Overall pass-through | Rent-sharing | Worker sorting

Firm observables (T4 col 3')

All firm controls (T4 col 5')

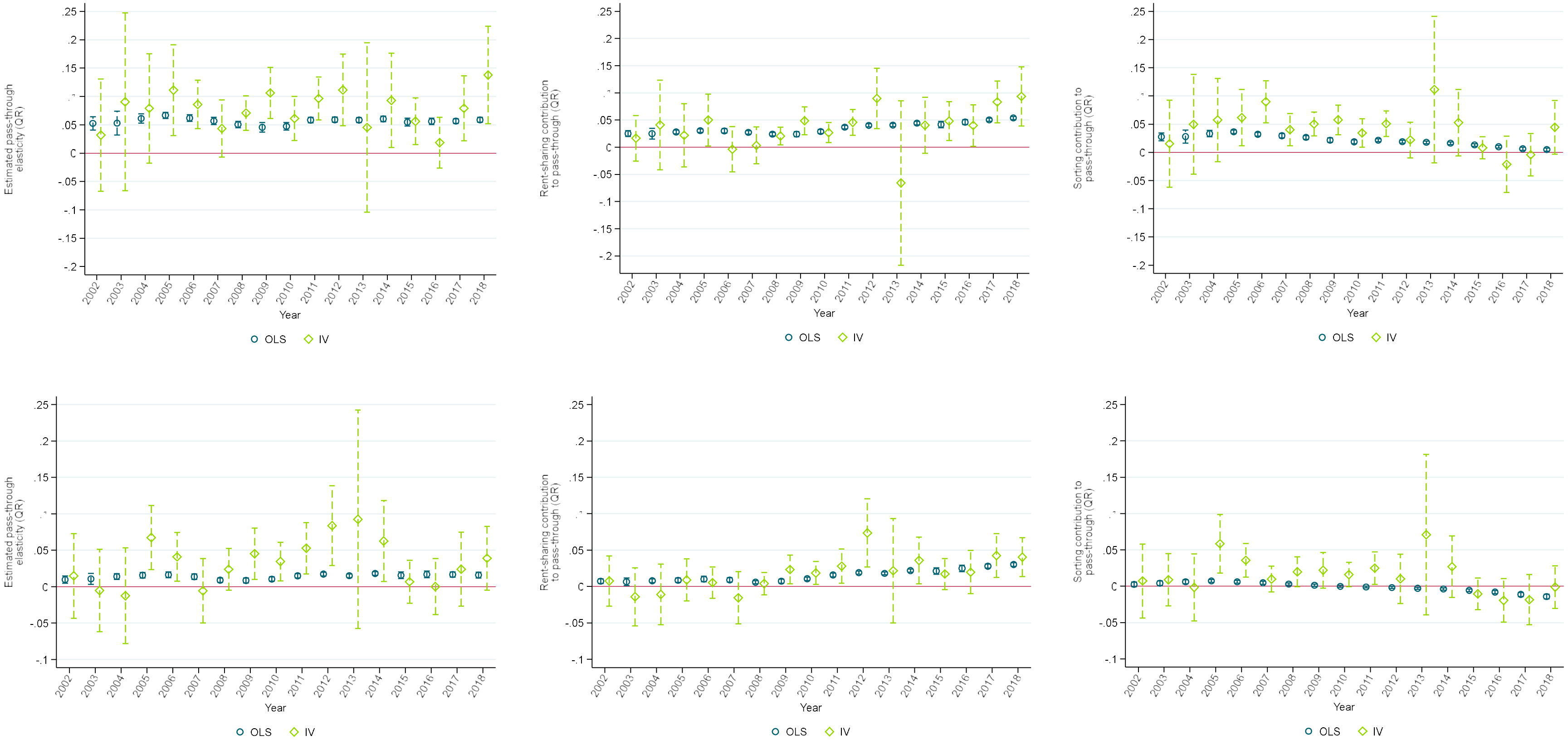


Figure B1 shows the time-varying estimates from our IV specification, with the OLS estimates plotted as a comparison. As with Figure 11, these are based on the firm observables specification (first row) and the firm fixed effects specification (second row). Results for overall pass-through are shown in the first column, rent-sharing in the second, and worker sorting in the third.

The first thing to notice is the difference in the width of the confidence intervals, which are orders of magnitude larger for the IV estimates than for the OLS estimates. IV estimates of overall pass-through (column 1) are generally higher, consistent with our baseline IV results. The IV estimates of overall pass-through are relatively stable over the period. We again see that the importance of rent-sharing in explaining pass-through has increased over the period, while the importance of worker sorting has declined. In the firm fixed effect specification, the contribution of worker sorting becomes negative at the end of the sample period, consistent with the OLS results. While the overall patterns in the IV point estimates are similar to the OLS estimates, the width of the confidence intervals make it difficult to draw strong conclusions about the changes in rent-sharing and worker sorting over time. The IV estimates include 17 endogenous variables and 49 instruments, making it a very large system of equations. The instruments are also rather weak, as shown in Table 4. Further work is required to deepen our understanding into the changing dynamics of worker sorting and rent-sharing.

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