Wages, employment, labour turnover, and the accessibility of local labour markets

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Abstract

In this paper, we extend a dynamic efficiency wage model to the case of multiple local labour markets that interact through migration. Firms are concerned about turnover costs. The quitting behaviour of workers is a function of local labour market conditions, non-wage income and the costs and benefits of migration to other local labour markets. A synthetic micro sample of 20,302 observations from the 1986, 1991 and 1996 New Zealand Censuses of Population and Dwellings provides evidence supporting the theory. Across subgroups, the wages of workers with relatively inelastic local labour supply and/or lower geographical mobility are relatively more responsive to changes in the local employment rate. The evidence is consistent with the notion that local employers engage in monopsonistic competition with respect to the employment of such workers.

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1. Introduction

During the 1990s some empirical phenomena in the labour market were documented that appeared to contradict the standard competitive model. An often-cited example is Card and Krueger's (1994) case study of the impact of a minimum wage increase, where, contrary to competitive theory, they found that an increase in the minimum wage increased employment in fast food outlets in New Jersey, United States. Another interesting case is the research of Blanchflower and Oswald (1990), who, using American and British data, found evidence for an inverse relationship between the level of pay of individuals and the prevailing local unemployment rate, which they labelled the 'wage curve'. Subsequently, they and many others have reported additional evidence for this relationship.[1]

A common thread running through such findings could be the presence of an upward-sloping supply curve facing the individual employer in a local labour market, rather than the perfectly elastic one of the competitive model. It is therefore no surprise that monopsony has gained a rather more prominent position in the theory of labour demand (see, e.g., Boal and Ransom, 1997; Manning, 2003). Nevertheless, outright monopsony seems a rather extreme market form, in that most firms face some competition in recruitment and barriers to entry into specific markets have been reduced rather than increased in recent decades. Thus, rather than assuming a single buyer of labour, Bhaskar and To (1999) formulated a theory of monopsonistic competition in which there is free entry, but with the establishment of new firms constrained by lump-sum start-up costs. Each employer then has some market power in the labour market, even though the firm employs only a small fraction of the work force. This wage setting power results from horizontal job differentiation, which is a form of worker heterogeneity that leads to workers preferring certain jobs on the basis

[1] Blanchflower and Oswald (1994) reported estimated wage curves for 12 countries: United States, United Kingdom, Canada, South Korea, Austria, Italy, Netherlands, Switzerland, Norway, Ireland, Australia and West Germany. Among others, see also Janssen and Konings (1998) on Belgium, Pannenberg and Schwartze (1998) on East Germany, Papps (2001) on New Zealand, Pekkarinen (2001) on Finland and Montuenga *et al*. (2003) for a comparison of wage curves from France, Italy, Portugal, Spain and the United Kingdom.

of non-wage characteristics. Monopsonistic competition may be responsible for the responses to unemployment or institutional shocks in local labour markets being quantitatively small.

One of the main reasons for horizontal job differentiation is the geography of local labour markets. Varying costs of job search and commuting create heterogeneity, even among otherwise identical workers, simply due to differing residential locations. If commuting costs are rather important, Bhaskar and To (1999, p. 195) find that employment will increase following a modest increase in the minimum wage.

The present paper develops this theme further by focussing on frictions in the labour market at a greater spatial scale than commuting, namely wage outcomes across well-defined local labour markets, linked through migration. Again heterogeneity of workers is introduced, but reservation wages now vary across workers because the lump-sum costs of migration between localities will depend on the location of their current job. The geography of labour markets then influences the relationship between local wages and employment. In addition, using a simple search model, Sato (2000) shows that as long as there are productivity differentials across local labour markets, those with higher productivity have higher equilibrium wages and lower unemployment rates. Even with low mobility costs, spatial real wage differentials can then persist in equilibrium because higher productivity regions have larger populations that result in utility-offsetting congestion costs (commuting costs and land rent).

In this paper we will carry out an empirical analysis of local wages and employment by means of data on a set of 30 urban labour markets in New Zealand. As with the models described above, the geography of labour markets and fixed costs of production are central features of our model and lead to monopsonistic tendencies in these local labour markets. We formulate a model that extends the approach of Campbell and Orszag (1998), who, in turn, combined elements of a dynamic efficiency wage model due to Phelps (1994) with endogenous quitting behaviour as in Salop (1979). We generalise quitting behaviour by including non-wage income and the possibility of migration to other local labour markets. Long-run inter-regional equilibrium is ensured by the equalisation of expected utility across local labour markets, as in Harris and Todaro (1970). Exogenous shocks that originate within the local labour market (such as a permanent change in the level of amenities) generate equilibria that lead to a positive long-run correlation between wages and unemployment, while shocks generated outside the local labour market lead to a negative correlation. Given an assumption of small local labour markets, the latter shocks are likely to dominate and generate Blanchflower and Oswald's wage curve, albeit with a varying elasticity in our model.

The paper extends previous research on the wage curve phenomenon in a number of ways. Firstly, the specification of our regression equations follows directly from the theoretical model, which enables us to recover parameters of the quit-rate function without having access to labour turnover data. Secondly, we introduce a spatial dimension, namely the accessibility of local labour markets, into estimation of the unemployment elasticity of pay. Thirdly, we control for spatial variation in real non-wage income, such as the variation in the purchasing power of national social security benefits. A fourth innovation in the present paper is that we link calculated wage curve elasticities for groups of workers to levels of geographical mobility observed for such workers in other research. For example, the extent of local monopsony captured by the wage curve appears to be greater for less geographically mobile groups, such as the lower skilled (e.g., Mauro and Spilimbergo, 1999).

Finally, we estimate the model proposed in this paper with data on New Zealand. This has several advantages. The first is that New Zealand has a small population of just over four million, with local labour markets dispersed over an area that is between that of Great Britain and Japan. While there is a mixture of large (such as Auckland and Wellington) and small urban labour markets, many of the local labour markets are 'thin' by international standards. The country provides therefore an interesting example of the case in which the geography of local labour markets may generate non-competitive influences on wage setting. Moreover, the distance between the 30 urban labour markets is such that inter-urban commuting would be virtually absent, so that there is no confounding of migration and commuting responses. A further interesting feature of the New Zealand case is that this country underwent a decade of drastic economic reform and restructuring since the mid 1990s. The impact of this has been well documented in the international literature (e.g., Evans et al., 1996). In the present context, of particular importance is legislation that was introduced in 1991 aimed at enhancing labour market flexibility (e.g., Morrison, 2003). By considering how the unemployment elasticity of pay changed over the decade with 1991 as midpoint, we can assess whether the reform may have had an impact on the unemployment elasticity of pay.

A 'synthetic' micro-level sample was generated from three New Zealand population censuses (1986, 1991 and 1996).[2] Each observation consists of a group of wage and salary earners at the local labour market level, defined according to their employment status, ethnicity, gender, age, education and occupation. For each group, an estimate of median annual income and mean weekly hours worked was obtained. This resulted in 20,302 useful observations. In addition, several local labour market characteristics were obtained for each urban area. Some of these were obtained from census data, some from other sources.

The next section sets out the theoretical model. Section 3 discusses the nature of the available data and provides some descriptive statistics. The empirical support for the model is discussed in Section 4 and Section 5 sums up.

2. A labour turnover cost model

Consider an economy with R local labour markets, indexed by $j = 1, 2, \ldots, R$. It is assumed that the number of local labour markets is relatively large and that no single market dominates the national economy. There are S identical firms in each local labour market, indexed by $i = 1, 2, \ldots, S$, with S being a large number. All firms produce the same homogeneous good, sold at a price of unity in all markets. Following Oi (1962), labour is considered a quasi-fixed factor of production, as the recruitment of new employees involves one-off training costs. Firms face a level of staff turnover that depends on the wage they offer relative to the reservation wages of workers and the alternative income available to them. Firms treat local average wages and employment levels as given. Given the assumption of identical firms, and an assumption of costless mobility within each small local labour market, wages and employment are equated across firms within local labour markets. However, due to the existence of location-fixed amenities and costly geographical labour mobility, wages may vary between local labour markets.

Firm i in labour market j aims to maximise discounted future profits at time t_0:

[2] The focus of this paper is in some ways similar to recent research by Bell *et al*. (2002), who estimate wage curves by means of British earnings data, motivated by a multi-regional labour market model with a long-run zero net migration condition. However, having time-series data for each region, Bell *et al*. are able to analyse regional dynamics and distinguish between Phillips curve and wage curve relationships (with evidence in favour of the latter). With our data limited to the pooling of three cross sections, we cannot explicitly investigate intra-regional wage dynamics in this paper.

$$\Pi_{ijt_0} = \int_{t=t_0}^{\infty} e^{-\rho t}((f(E_{ijt}) - w_{ijt}E_{ijt} - T(h_{ijt})E_{ijt})\mathrm{d}t\,, \tag{1}$$

subject to the dynamic constraint:

$$\frac{\dot{E}_{ijt}}{E_{ijt}} = h_{ijt} - q_{ijt}\,. \tag{2}$$

Here E_{ijt} represents the firm's employment level, $f(E_{ijt})$ the production function, ρ the discount rate, w_{ijt} the wage, h_{ijt} the firm's hiring rate, $T(h_{ijt})$ the training cost and q_{ijt} the endogenous quit rate, to be specified below. The training cost, $T(h_{ijt})$, is measured in terms of the cost to the firm per existing employee of training new recruits, who arrive at rate h_{ijt}.[3] The price of output has been set as the numeraire, therefore wages can be interpreted as real values.

The current-value Hamiltonian for firm i in region j is as follows, where λ_{ijt} is the Lagrange multiplier:

$$H_{ijt} = f(E_{ijt}) - w_{ijt}E_{ijt} - T(h_{ijt})E_{ijt} + \lambda_{ijt}(h_{ijt} - q_{ijt})E_{ijt}\,. \tag{3}$$

The first-order conditions for this maximisation problem are:

$$T'(h_{ijt}) = \lambda_{ijt}\,; \tag{4}$$

$$-\lambda_{ijt}\frac{\mathrm{d}q_{ijt}}{\mathrm{d}w_{ijt}} = 1\,; \tag{5}$$

$$-\frac{\partial H_{ijt}}{\partial E_{ijt}} = \dot{\lambda}_{ijt} - \rho\lambda_{ijt}\,. \tag{6}$$

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[3] Since it is paid for by firms, training is implicitly assumed to be job-specific, along the lines of Becker (1962).

The interpretation of these first-order conditions is straightforward. Equation (4) equates the marginal benefit from recruiting an additional worker, λ_{ijt}, to the marginal training cost. Equation (5) says that the marginal benefit from increasing the wage, in terms of reduced quits, should equal the marginal cost of the rising wage bill. Equation (6) expresses the marginal benefit of recruiting an additional worker in terms of the change in the present value of discounted future profits.

Following Campbell and Orszag (1998) and others, the training cost function is assumed to be a quadratic:

$$T(h_{ijt}) = \frac{A_t}{2}(h_{ijt})^2, \qquad (7)$$

in which A_t is a potentially time-variant parameter. With respect to the quit rate, it is assumed, along the lines of Phelps (1994), that quits are a function of the ratio of the wage offered by the firm to the alternative wage offered by other firms in the local labour market. Specifically, quits are inversely related to $\dfrac{w_{ijt}}{(1-\pi(u_{jt}))w_{\cdot jt}}$, where $w_{\cdot jt}$ is the average wage in the local labour market and $\pi(u_{jt})$ is the instantaneous probability of becoming unemployed after quitting. The latter is unobservable but is assumed to be increasing in the observed unemployment rate in the local labour market, u_{jt}. Next, we assume that quitting is also inversely related to the ratio of the wage to local non-market income (such as a social security benefit), $\dfrac{w_{ijt}}{b_{jt}}$, where b_{jt} is the local non-market income.

Like most efficiency wage models, the model proposed by Campbell and Orszag (1998) describes a single-point economy. We now extend this model to an explicit spatial setting by allowing for the possibility of migration. Thus, we assume that quits are affected by the wages available in other local labour markets. Because individual local labour markets are assumed to exert only a small influence on the national level aggregates, we can approximate the unemployment rates and wage levels of other local labour markets (that is, labour markets $r \neq j$) by the economy-wide values. We assume therefore that quits are inversely related to $\dfrac{w_{ijt}}{\gamma(c_{jt})(1-\pi(u^e_{\cdot t}))w^e_{\cdot\cdot t}}$, where $w^e_{\cdot\cdot t}$ is the expected economy-wide wage at time t and $\pi(u^e_{\cdot t})$ is the expected economy-wide probability of being unemployed after quitting at time t. By definition, $w^e_{\cdot\cdot t}$ and $u^e_{\cdot t}$ are the weighted averages of each local labour market's expected wage and expected unemployment rate, weighted by local employment and the labour force, respectively. $\gamma(c_{jt})$ adjusts the expected economy-wide average wage downward to take account of the average migration cost, c_{jt}, incurred in migration from labour market j, hence it is a weighted average of migration costs with respect to all other local labour markets. The function $\gamma(\cdot)$ has $\gamma' < 0$, $\gamma'' > 0$ and, in the case of costless mobility, $\gamma(0) = 1$.

The quit-rate function is assumed to be of constant elasticity with respect to its arguments and, given the above assumptions, may be written as:[4]

$$q_{ijt} = B_t \left(\frac{w_{ijt}}{(1-\pi(u_{jt}))w_{\cdot jt}} \right)^{-\eta} \left(\frac{w_{ijt}}{b_{jt}} \right)^{-\nu} \left(\frac{w_{ijt}}{\gamma(c_{jt})(1-\pi(u^e_{\cdot t}))w^e_{\cdot\cdot t}} \right)^{-\mu}, \tag{8}$$

in which B_t is a potentially time-variant parameter. Substituting the training cost function (7) and the quit-rate function (8) into the first-order conditions (4) and (5) yields the following equations:

$$A_t h_{ijt} = \lambda_{ijt}; \tag{9}$$

$$\lambda_{ijt} B_t (\eta + \nu + \mu) w_{ijt}^{-\eta-\nu-\mu-1} ((1-\pi(u_{jt}))w_{\cdot jt})^{\eta} b_{jt}^{\nu} (\gamma(c_{jt})(1-\pi(u^e_{\cdot t}))w^e_{\cdot\cdot t})^{\mu} = 1. \tag{10}$$

We now focus on the steady-state equilibrium. Given the assumptions of the model, all firms within a local labour market offer the steady-state wage. Employment in each firm is also constant. Consequently, this also extends to the economy-wide expected wage and employment levels. Furthermore, we take the local non-market income and migration cost as given. Hence, the following conditions must hold for i = 1, 2, …, S and j = 1, 2, …, R:

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[4] Interaction effects between the three terms in parentheses are assumed absent, but the chosen multiplicative form has the advantage of generating additive components in a log-linear earnings function.

$$h_{ijt} = q_{ijt}\,; \tag{11}$$

$$w_{ijt} = w_{.j\cdot}\,;\; u_{jt} = u_{j\cdot}\,;\; w^{e}_{..t} = w^{e}_{...}\,;\; u^{e}_{.t} = u^{e}_{..}\,;\; b_{jt} = b_{j\cdot}\,;\; c_{jt} = c_{j\cdot}\,;\; A_t = A_{.}\,;\; B_t = B_{.} \tag{12}$$

Combining Equations (9) to (12) then yields the following relationship involving the steady-state wage level, $w_{.j.}$, and unemployment rate, $u_{j.}$, in region j:

$$A_{.}B_{.}^{2}(\eta+\nu+\mu)(1-\pi(u_{j.}))^{2\eta}\,w_{.j.}^{-2\nu-2\mu-1}b_{j.}^{2\nu}(\gamma(c_{j.}))^{2\mu}(1-\pi(u^{e}_{..}))^{2\mu}(w^{e}_{...})^{2\mu} = 1\,. \tag{13}$$

Taking the natural logarithm of both sides and simplifying by the introduction of some new parameters allows Equation (13) to be rewritten as follows:

$$\ln w_{.j.} = \kappa + \xi\ln b_{j.} + \varphi\ln(\gamma(c_{j.})) + \delta\ln(1-\pi(u_{j.}))\,, \tag{14}$$

where:

$$\kappa = \frac{\ln(\eta+\nu+\mu) + \ln A_{.} + 2\ln B_{.} + 2\mu(\ln(1-\pi(u^{e}_{..})) + \ln w^{e}_{...})}{2\nu+2\mu+1}\,; \tag{15}$$

$$\xi = \frac{2\nu}{2\nu+2\mu+1}\,; \tag{16}$$

$$\varphi = \frac{2\mu}{2\nu+2\mu+1}\,; \tag{17}$$

$$\delta = \frac{2\eta}{2\nu+2\mu+1}\,. \tag{18}$$

Equation (14) shows that there is a log-linear relationship between the local wage level, $w_{.j.}$, and the probability of re-employment after quitting, $1-\pi(u_{j.})$. We will refer to this as the wage curve (*WC*) relationship, as it defines the local wage level, $w_{.j.}$, to be a monotonically declining function of the local unemployment rate, $u_{j.}$. Because of the assumption of small local labour markets, changes in any given local labour market do not affect expectations of national wages and unemployment.

Consequently, a positive productivity shock in the local labour market would increase the real wage and decrease unemployment, and *vice versa* with a negative employment shock. In isolation, the local labour market exhibits a wage curve over the business cycle.

In long-run spatial equilibrium, all local labour markets must offer a given worker the same level of expected utility. If this were not the case, migration would take place until either some regions were completely depopulated of workers or all inter-regional utility differences were eliminated. Hence, it shall be assumed that the following spatial equilibrium condition is satisfied for all $j = 1, 2, \ldots, R$:[5]

$$(1-\pi(u^e_{j.}))w^e_{.j.} = \zeta_j (1-\pi(u^e_{..}))w^e_{...}\,. \tag{19}$$

The parameter ζ_j allows for deviations from the assumptions of perfect mobility and identical regional attributes. If $\zeta_j = 1$ for all j, Equation (19) is equivalent to the standard Harris and Todaro (1970) condition that expected wages are equalised across regions. However, there are two reasons why ζ_j may differ from unity: the existence of location-fixed amenities that affect the utility of workers and imperfect arbitrage due to the pecuniary and non-pecuniary costs of migration.

We will now assume that fluctuations in wages and unemployment are stationary about a constant mean, which is therefore the optimal long-run forecast for each variable. Hence, the steady-state values are used in place of the expected values in (19). A rational expectations spatial equilibrium is then defined by (14) jointly with the following condition:

$$\ln w_{.j.} = \ln(\zeta_j) + \ln(1-\pi(u_{..})) + \ln w_{...} - \ln(1-\pi(u_{j.}))\,, \tag{20}$$

which follows directly from (19) by taking logarithms and substituting the steady-state values for the variables. Given its origin, Equation (20) will be referred to as the Harris and Todaro (*HT*) condition. The local steady-state relationship *WC* and the long-run spatial equilibrium condition *HT* are depicted in Figure 1. The vertical axis

[5] Equation (19) is the outcome of the workers' utility maximisation problem that need not be explicitly formulated here since, in addition to the equilibrium condition, worker behaviour is fully specified by the quit-rate function (8).

corresponds to zero unemployment. The long-run spatial equilibrium is determined by the intersections of *R* pairs of *WC* and *HT* lines.[6]

The *HT* lines have a slope of one. The position of the *HT* lines varies across local labour markets with spatial variation in local amenities. A decrease in amenities is likely to lead to a compensating increase in the long-run wage level and a decrease in the long-run unemployment rate. As long as the impact of cross-section variation in non-wage income and migration costs (given by $\frac{2\nu}{2\nu+2\mu+1}$ and $\frac{2\mu}{2\nu+2\mu+1}$, respectively) is small, we are likely to find that in a cross-section of local labour markets, the scatter diagram of wages and unemployment rates shows an inverse relationship, *i.e*. a wage curve. As in Blanchflower and Oswald's (1994) labour contract and efficiency wage models, regions with unattractive amenities have high (compensating) wages and low unemployment (i.e. they are in the north-west of Figure 1) and regions with attractive amenities are in the south-east.

FIGURE 1 ABOUT HERE

Over time, permanent shocks in $w_{\cdot\cdot\cdot}$ and/or $u_{\cdot\cdot}$ also generate an inverse relationship between local wages and unemployment, as such shocks shift the *WC* line less than the *HT* line. Hence, when there is a permanent increase in national demand, the equilibrium wage will increase and the unemployment rate will decrease in all local labour markets. On the other hand, a permanent change in the available non-wage income or accessibility of a particular local labour market would move the equilibrium along the corresponding *HT* line, *ceteris paribus*. Hence, permanent shocks that originate *within* the local labour market generate equilibria that lead to a positive correlation between wages and unemployment, while permanent shocks generated outside the local labour market lead to a negative correlation. Because of the assumption of small local labour markets, the latter shocks are likely to dominate the former. After describing the data in the next section, we will discuss estimation of the *WC* lines in Section 4.

[6] Of course, $u_{j\cdot}$ and $w_{\cdot j\cdot}$ are included in the calculation of the national averages $u_{\cdot\cdot}$ and $w_{\cdot\cdot\cdot}$, which are part of the intercept of the *HT* and *WC* lines. However, because the number of local labour markets, *R,* is large, each has a negligible impact on the national averages, which allows us to draw Figure 1. However, even if we do not make that assumption, the system of *2R* equations in the *2R* variables $u_{j\cdot}$ and $w_{\cdot j\cdot}$ ($j = 1,\ldots,R$) can be solved iteratively and the figure drawn subsequently.

3. The New Zealand dataset

The theory of the previous section was tested using data from the 1986, 1991 and 1996 New Zealand Censuses of Population and Dwellings. As noted in the introduction, the geography of local urban labour markets in New Zealand is such that this country would provide a particularly appropriate case to test the proposed theory. Estimation of Equation (14) requires local labour market data on real wages, employment, unemployment, real non-wage income and the cost of migration. We constructed a 'synthetic' individual-level sample of workers aged between 15 and 59. The use of this dataset was motivated by the limited availability of unit record data in New Zealand in terms of access and cost, due to legislation protecting the confidentiality of the original records. Public use samples of censuses or large surveys, that are readily available in some other countries, are quite rare in New Zealand. However, recent initiatives by Statistics New Zealand, the government agency responsible for official statistics, will enhance the access to 'confidentialised' unit records. This will enable re-estimation of the proposed model with unit record data at a later time.

Observations on median annual income, Y, and average weekly hours normally worked, H, were obtained for 34,560 population subgroups of wage and salary earners aged between 15 and 59. Non-wage and salary earners (for example, self-employed individuals) were excluded as the earnings of these individuals are likely to be determined by a process that does not conform to the model presented in the previous section. The subgroups form a mutually exclusive and exhaustive set of permutations of the following variables: year (3 values); urban area (30 values); age (3 values); educational attainment (2 values); ethnicity (2 values); gender (2 values); full-time/part-time status (2 values); occupation (8 values).[7] The 30 urban areas are assumed to constitute single local labour markets. As distances between these are large, there is virtually no inter-urban commuting.

Attention was restricted to workers in 30 urban areas and 8 occupations in order to control for as many unobservable characteristics as possible, in a sense allowing each group to closely 'resemble' an individual worker. A drawback of this approach is that

[7] Details of these groups are listed in the Appendix. Each occupational group comprises a number of related four-digit occupations.

it will lead to selectivity bias in the regressions performed in the next section if workers in either omitted locations or omitted occupations have different unobservable characteristics to included workers. Although the latter is probable, the groups are broadly representative of the labour market, although not always in the same proportions as in the national data. Nonetheless, any conclusions drawn should only be interpreted as applying to workers in the specific occupations and urban areas chosen.

Certain combinations of characteristics yielded zero frequencies. These observations were deleted, leaving 20,302 observations in total. Collectively, around half a million workers are covered in each year, or approximately 56% of all wage and salary earners in New Zealand.[8]

The wage, w, featured in the model presented in the previous section represents the real hourly earnings of workers. The census data only provided information on annual income and weekly hours. The relationship between the theoretical wage and the available information is as follows:

$$w_{gjt} = \frac{\theta_{gjt} Y_{gjt}}{52\lambda_{gjt}\left(1-\widetilde{u}_{gjt}\right)H_{gjt}P_t}, \tag{21}$$

where w_{gjt} is the theoretical hourly wage in constant 1986 dollars of a group g in urban area j at year t. This is calculated by dividing annual income Y_{gjt}, times the fraction of income derived from salary and wages θ_{gjt}, by the national Consumers' Price Index value P_t and the actual hours worked per year. The latter are the product of the usual hours worked per week H_{gjt} times the number of weeks worked. The number of weeks worked is given by the group's fraction of 52 weeks participating in the labour market λ_{gjt} and the fraction of those weeks in employment (1-$\widetilde{u}_{gjt}$).[9] Following Card (1995), we call $\widetilde{u}_{gjt}$ the person's retrospective unemployment rate. With cross-sectional grouped data from three censuses, θ_{gjt}, λ_{gjt} and $\widetilde{u}_{gjt}$ are

[8] 532,755 wage and salary earners are included in the 1986 sample, 497,136 in 1991 and 565,788 in 1996.

[9] Equation (21) assumes that H is the average hours worked per week over the previous 52 weeks, which differs from the census definition in 1986 and 1991, when respondents were asked their hours worked in the previous week. Consequently, for these years, the predicted values from a regression of reported hours on the 1996 hours for each group were used, respectively.

unobserved, but the sample of individuals for whom census data were obtained was restricted to salary and wage earners employed at the time of the census. Most of these were in the labour force throughout the year and had experienced no spell of unemployment in the previous twelve months. For example, New Zealand longitudinal data showed that typically less than ten percent of working age population experienced unemployment at least once during a given year (e.g., Gobbi and Rae, 2000). Similarly, gross worker flow data show that the inflow rate from out of the labour force to employment is less than 5 percent per quarter (e.g. Silverstone and Gorbey, 1995). Furthermore, non-wage income among the employed 15-59 year olds is typically small and would not have varied much across the selected groups. Consequently, for a large proportion of the 20,302 observations, $\theta_{gjt} \cong 1$, $\lambda_{gjt} \cong 1$ and $\tilde{u}_{gjt} \cong 0$. We therefore proxy the theoretical real wage by

$$\hat{w}_{gjt} = \frac{Y_{gjt}}{52 H_{gjt} P_t}, \tag{22}$$

and, given the selected groups of employed salary and wage earners, the effect of this approximation is not expected to influence the statistical inference.[10]

A variety of aggregate statistics for each urban area were collected from the 1981, 1986, 1991 and 1996 Censuses, with the 1981 data being included to permit lagged variables in the model at all three subsequent census dates. These include the working age population, N, the number of unemployed males and females and the number of each gender in the labour force. The variables u_T, u_M and u_F refer to the total, male and female unemployment rates, while the variables L_T, L_M and L_F refer to the respective participation rates. Labour market slackness in each local labour market is measured by the gender-specific unemployment rate.

[10] For varying λ_{gjt} and $\tilde{u}_{gjt}$, the relationship between the 'true' elasticity of interest, δ, and the estimated elasticity by means of $\hat{w}_{gjt}$ is then as follows: $\delta = \frac{d\ln w}{d\ln(1-u)} = \frac{d\ln \hat{w}}{d\ln(1-u)} - \frac{d\ln\lambda}{d\ln(1-u)} - \frac{d\ln(1-\tilde{u})}{d\ln(1-u)}$ in which $\frac{d\ln \hat{w}}{d\ln(1-u)}$ is the calculated elasticity and u is the observed unemployment rate in the local labour market. New Zealand evidence confirms the international consensus that the discouraged worker effect tends to outweigh the added worker effect (e.g. Morrison, 1999) and the retrospective risk of unemployment of the individual would be positively correlated with the local unemployment rate but, averaged out over the 20,302 groups of employed salary and wage earners chosen, the latter two elasticities are expected to be rather small. We conclude that available data generate a slight, but unquantifiable upward bias in estimation of the unemployment elasticity of pay.

In the absence of data on local cost-of-living variation, which determines the real value of non-wage income, we used the only readily available indicator, namely the average price of houses sold in each urban area. This information was derived from biannual urban property sales statistics.[11] Each average price was expressed as a proportion of the corresponding national value for the particular year in order to remove the effect of overall growth in house prices in New Zealand. This relative house price, p_{jt}, therefore represents the extent of cross-sectional variation in house prices in a given year.[12]

A plausible proxy for the cost of migration is the average time it takes migrants from urban area j to reach other urban areas by road. This involved estimating the road travel time in hours between pairs of urban areas.[13] The average travel time to other local labour markets from urban area j at time t, d_{jt}, was then calculated by weighting each travel time by the size of the labour force in the urban area of destination (thus taking account of the volume of job opportunities there).[14] Hence, the weighted average travel time from urban area j is determined as follows, where T_{jk} is the travel time between urban areas j and k:

$$d_{jt} = \frac{\sum_{k=1}^{30} T_{jk} L_{kt} N_{kt}}{\sum_{k=1}^{30} L_{kt} N_{kt}}, \tag{23}$$

and L and N were defined earlier. Table 1a presents the mean and standard deviation across subgroups for each variable by census year. These are calculated by weighting observations by the frequency in each cell. This implies that the reported values reflect the average characteristics of all workers in the sample. For comparison

[11] The average values obtained were for the territorial authority, or authorities, in which each urban area is located. While the boundaries of these underwent substantial changes in 1989, this did not affect noticeably the cross-sectional variation in housing costs in our data set.

[12] Hence, p_{jt} is distinct from P_t, which is a time series of national Consumers' Price Index values.

[13] Where two urban areas are located in different islands this includes the time taken to cross Cook Strait by car ferry.

[14] Travel time is a better proxy for migration costs than, say, distance by road as it captures the effect of physical barriers, such as the Southern Alps and Cook Strait. d is not time-invariant as it is affected by changes in the spatial distribution of the labour force over time.

purposes, corresponding statistics for the entire population of New Zealand wage and salary earners are reported in Table 1b. The sample contains more females and part-time workers than the labour force as a whole. As noted earlier, this implies that conclusions drawn at the national level should be interpreted as simply applying to the occupations and urban areas chosen. Means and standard deviations for each urban area-level variable are given in Table 2. Here observations are weighted by the working-age population of the particular urban area.

TABLE 1 ABOUT HERE

TABLE 2 ABOUT HERE

Not surprisingly, Table 1a indicates that average nominal annual income, Y, increased over the sample period, albeit less rapidly between 1991 and 1996. Comparing these values with the price level, P, reveals that real income peaked in 1991, with price inflation exceeding wage growth over the ensuing five years.[15] The decline in real income in New Zealand at national and regional levels during a period of labour market reform has been well documented (e.g., Karagedikli et al., 2000). Average hours worked, H, also declined, largely reflecting a corresponding fall in the proportion of full-time workers in the labour force. The trend in the average estimated real wage, $\hat{w}$, over the decade reflected changes in both real income and hours worked. Wages rose slightly between 1986 and 1991, but then fell sharply, so that workers sampled in 1996 earned less per hour on average than those sampled ten years earlier.

The 1991 Census coincided with a period of recession, as noted by the peak in the unemployment rate in the sample. The sample from that year is also the 'oldest', with only 29% of workers in the 15-25 age group, and the most educated, with 44% having a post-secondary school qualification. The absence of younger and less skilled individuals from the sample would appear to support the notion that workers with less job-specific training are more likely to be laid off during periods of recession.

In general, Table 1 provides evidence of a shift from low-skilled to high-skilled occupations. The proportion of the sample and of the labour force employed as teachers, salespeople or service workers rose from 1986 to 1996, although the

[15] P takes the value 1 for 1986, 1.499 for 1991 and 1.643 for 1996.

proportion of health professionals peaked in 1991. On the other hand, the proportions of production workers and construction workers in the sample fell over time, partly due to a net outward migration of these types of workers to Australia (Morrison, 2001). An exception to this pattern is the employment of labourers, who comprised a greater proportion of the 1996 sample than that of ten years earlier, despite being a low-skilled occupation. Consistent with international trends, the proportion of women in the sample increased over the decade.[16]

In addition to the above variables, a set of additional variables was collected to be used as instruments for the endogenous determinants of wages. The proportion of private dwellings that were rented in each urban area, r, was obtained from the census. Data on annual rainfall and sunshine hours by urban area and year were also collected and were denoted W_R and W_S, respectively. Lastly, an industry-mix variable, I, was constructed along the lines of Bartik (1991). This is a weighted average of the growth rates of national industry employment over the year prior to each census, with the weights calculated as the share of a local labour market's employment in each industry at the time of the census. In other words, I would represent the employment growth rate for each urban area if all sectors grew at the national rate. Means and standard deviations for these four variables are reported in Table 2.

4. Empirical analysis

In this section we estimate the *WC* line derived in Section 2 by means of New Zealand data. The *WC* line represents a locus of intra-local labour market steady states, while the *HT* line characterises the condition for long-run spatial equilibrium. In the short run, local labour market outcomes may be on the *WC* line but off the *HT* line. By comparing long-run equilibria, Papps (2001) established evidence of a positive relationship between the long-run local wage and estimates of the local rate of 'permanent' unemployment. Hence, these long-run equilibria are on the *HT* line, as expected.

[16] Among the occupations chosen for this study women make up a disproportionately large share of workers. Nonetheless, as shown in Table 2, a similar trend occurred in the total workforce, with the fraction of males falling from 57% in 1986 to 51% in 1996.

The turnover cost model outlined in Section 2 predicts that the wage in a local labour market is a function of the probability of re-employment after quitting, $1-\pi(u)$. We approximate this probability by one minus the unemployment rate, which we refer to as the employment rate.

The model of Section 2 assumes that both workers and jobs are homogeneous, but any empirical estimates of Equation (14) must also include workers' human capital variables and job characteristics as regressors to allow for the existence of a heterogeneous workforce in reality. Accordingly, dummy variables are added to control for a group's age, educational level, ethnicity, gender and employment status. To account for exogenous productivity growth over time, time fixed effects are included. Inter-urban variation in the intercept of (14), which can be due to inter-urban differences in training costs, quitting behaviour or expectations formation, are captured by spatial fixed effects. Furthermore, it is assumed that nominal non-wage income is identical across local labour market areas. This is a reasonable assumption as there is little variation in nominal social security benefits across New Zealand's local labour markets. Thus, variation in real non-wage income is measured by the reciprocal of the local cost-of-living index. The latter is approximated by the relative house price index, p_{jt}.[17]

In the model, the remaining determinant of the wage is the average migration cost from local labour market j. The effect of this is given by $\varphi \ln(\gamma(c_{jt}))$ in Equation (14). Since this factor is not directly observable, the negative of the weighted average travel time by road, $-d_{jt}$, as defined in the previous section, is used as a proxy for $\ln(\gamma(c_{jt}))$. This reflects the relative difficulty with which workers can migrate in and out of labour market j, *i.e.* the 'remoteness' of the local labour market in the system.

The estimable equivalent of equation (14) is therefore as follows:

$$\ln w_{gjt} = \kappa^* + \omega_j + \iota_t + \mathbf{x}_{gjt}'\boldsymbol{\gamma} + \xi \ln p_{jt}^{-1} + \varphi(-d_{jt}) + \delta \ln(1-u_{gjt}), \tag{24}$$

[17] Our interpretation of spatial house price variation is different from that of Bell *et al*. (2002) who consider this a proxy of (dis)amenities. Inter-urban house price variation is primarily due to the impact of agglomeration effects on productivity and land prices. For example, house prices are on average relatively high in the large cities. Utility-yielding urban amenities could be offset by disutility-yielding congestion and crime. Instead, there is some evidence of beneficiaries moving out of the larger urban areas in order to increase the purchasing power of the nominally fixed social security benefit, which is consistent with our real non-wage income interpretation of p_{jt}^{-1} (Morrison and Waldegrave, 2002).

in which ω_j and ι_t are the location and time fixed effects respectively and $\mathbf{x}_{gjt}$ is a vector of personal characteristics of the salary and wage workers.

Estimates of Equation (24) are reported in Table 3. Column (1) reports the results of a feasible generalised least squares (GLS) regression. This method corrects the problem identified by Moulton (1990) that estimation of (24) with OLS underestimates standard errors of regressors which vary only across gender and location, but not across groups. We use the White (1984) correction with clustering on region/gender groups. This method allows for arbitrary intra-cluster correlations and Monte Carlo experiments suggest that the method performs well against other ways of correcting for the Moulton problem (see Bertrand et al., 2004). In addition, each observation is weighted by the number of persons in the underlying group.

Column (1) of Table 4 shows that the employment rate is found to have a significantly positive impact on wages, as predicted by the model. The estimated employment elasticity of pay, $\hat{\delta}$, is 1.894, meaning that a 10% increase in the employment rate of a local labour market coincides with an 18.9% increase in wages. All demographic variables enter the regression equation with the expected coefficients. Apart from the coefficient on the dummy variable for full-time workers, which is significant at the 10% level, these are all significant at the 1% level.

TABLE 3 ABOUT HERE

The level of real non-wage income, proxied by $\ln p^{-1}$, has a positive impact on the wage, as expected, although significant only at a 22% level. This weak result is undoubtedly related to the absence of accurate local cost of living data and the use of the cost of housing as a proxy, as noted earlier in the paper.

The coefficient of the measure of accessibility, $-d$, is also positive, as expected, and is significant at the 1% level. This confirms the prediction of our turnover cost model that, *ceteris paribus*, relatively isolated urban areas feature lower wages. Lower migration costs will increase the propensity of workers to quit and, hence, increase the wage they are paid.

Because the employment rate in long-run spatial equilibrium is determined by the intersection of the *WC* and *HT* lines in the model of Section 2, the above estimates of δ may be biased and instrumental variables (*IV*) estimation may be needed. A

Hausman test suggested that ln(1–u) is indeed an endogenous variable and therefore requires instrumentation. In addition to the four contemporaneous instruments defined in the previous section (*viz*. housing tenure, rainfall, sunshine and industry mix), the corresponding values of ln(1–u) from the previous census are also used for this task.

Table 4 provides evidence of the suitability of these variables to serve as instruments. All are correlated with the employment rate at the 1% level. However, none are significantly correlated with the residuals, ε, from the regression underlying the first column of Table 3, suggesting that they are appropriate instruments.

TABLE 4 ABOUT HERE

Nonetheless, finding good instruments is difficult and particularly the low correlation between ln(1–u) and its value at the previous census five years earlier is noticeable. The second column of Table 3 treats ln(1–u) as endogenous and reports the results of IV estimation using the set of instruments discussed above. The demographic variables remain statistically significant, however the estimated coefficient on the employment rate, $\hat{\delta}$, falls to 0.822 and is no longer significant. The coefficient on the labour market accessibility variable is now also insignificant. These less satisfactory results may be related to the limited availability of suitable instruments.

What do these regressions imply with respect to the wage curve elasticity? Because the employment rate, rather than the unemployment rate, is included in the earnings equation, the unemployment elasticity of pay is not constant. Instead, it can be computed from:

$$\frac{\partial \ln w_{gjt}}{\partial \ln u_{gjt}} = \frac{u_{gjt}}{u_{gjt} - 1}\delta . \tag{25}$$

At the mean, the preceding regressions imply a value of β that lies between –0.092 (0.157) with instrumental variables to –0.213 (0.020) with GLS (standard errors in parentheses). The latter is somewhat greater than the ‘consensus’ wage curve elasticity of –0.1, and may be slightly biased upwards due to the wage measurement problem referred to in the previous section, but within the range of estimates found by others (see Nijkamp and Poot, 2002).

The estimated coefficients in columns (1) and (2) of Table 4 were constrained to be identical across genders. Many wage curve studies find an unemployment elasticity of pay that is greater for males than for females (see, e.g., Card, 1995; Baltagi and Blien, 1998; Janssens and Konings, 1998). This could be due to the more elastic labour supply of the latter (e.g., Killingsworth,1983). To test for this effect with the New Zealand data, we re-estimate the model with the IV method for males and females separately. The results are in columns (3) and (4) of Table 3 respectively.

Some well-known results are found. For example, seniority (as measured by age) increases the earnings of men more than those of women. Men aged 41-60 earn 53% more than those aged 15-25, while for women the corresponding value is 33%. The earnings premium for post-secondary school qualifications is greater for men than for women. The employment elasticity is positive and statistically significant for men, but negative and insignificant for women, which is consistent with the evidence from several other countries noted above. The estimated coefficient for men implies an unemployment elasticity of pay of –0.217 at the mean. The local labour market accessibility effect is positive for males, but only significant at a 14% level. On the whole, the female equation does not support the model embodied in Equation (24), while the male equation is reasonably convincing.

Equation (18) suggests another reason why the employment elasticity of pay may vary across groups of workers. Workers with low values of η and high values of μ will also have lower employment elasticities. Geographically mobile workers may be relatively less responsive to changes in local labour market conditions and more responsive to changes in conditions elsewhere. Thus such workers may have low values of η and high values of μ. The argument is consistent with Topel's (1986) claim that 'the wage and unemployment consequences of within-area changes in labour demand fall on those with the strongest area attachments, that is, those who are least mobile in response to current and expected area wage differentials' (p. S141).

Hence, a steeper *WC* line should be found for less geographically mobile groups of workers. This hypothesis was tested by means of re-estimation of equation (24) for sub-groups. In order to control for the Moulton (1990) problem, the reported estimates were obtained by means of the feasible GLS method.[18] These can be found in Table 5. For comparison, the estimate of δ from column (1) of Table 3 is again reported at the top of the $\ln \hat{w}$ column of Table 5.

TABLE 5 ABOUT HERE

As before, the employment elasticity is significant for males, but not for females. The level calculated with the GLS method, however, is rather smaller than with the IV method and implies a wage curve elasticity of -0.04 for males. Another interesting issue is the wage responsiveness across age groups and education. Topel (1986) identified older workers and those with less education as being the least mobile workers and, hence, the most vulnerable to changes in local labour market conditions. The results presented in Table 5 provide some support. Changes in the employment rate are found to influence the wages of workers with no post-secondary school qualification more than the wages of those with a qualification, in accordance with Topel's prediction (see also Johansen, 1999). However, the employment elasticity of pay has an inverted U-shaped relationship with age, which in our model setting is consistent with lower geographical mobility of the middle age group (possibly resulting from the greater cost of migration of households with children and a lesser wage elasticity of labour force participation).

Theory provides no clear prediction of whether or how the wage curve will vary across occupations. However, workers with skills that are more easily transferred to another labour market should be less vulnerable to local business conditions. Using aggregate New Zealand data, Morrison and Poot (1999) found evidence that the wage curve is more elastic for relatively less-skilled occupations. Table 5 suggests that office clerks, service workers and production workers have the highest employment elasticities of pay. This would seem to confirm our expectation that local monopsony power has a stronger impact on the pay of those workers than on those of the higher skilled. No wage curve is found for the relatively geographically mobile groups, such as health professionals, teachers, salespeople or construction workers for whom the labour market across the Tasman Sea in Australia is an important source of employment opportunities.

Part-time workers appear to have a much larger employment elasticity of pay than full-time workers, suggesting that the former may be more likely to be what are referred to as 'tied stayers' in the migration literature, that is, less mobile and attached to the local labour market due to the employment of a partner.

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[18] IV estimates for sub-groups can be obtained from the authors upon request.

Table 5 also suggests that there are important differences in the responsiveness of wages to changes in local employment conditions for workers of different ethnicities. Even after controlling for all available measures of human capital, the logarithm of the employment rate is found to enter the earnings equation of those of European descent with a significantly smaller coefficient than is the case for non-Europeans. This suggests that those of non-European descent (primarily New Zealand Māori and Pacific Islanders) are more exposed to local monopsony power than Europeans, which is quite plausible given the disadvantaged position of the former in New Zealand labour markets (e.g., Herzog, 1997).

Table 5 also reports the employment elasticity of pay for the three census years separately. As noted in the introductory section, the 1986-96 decade was marked in New Zealand by significant liberalisation, reform and adjustment in all sectors of the economy. Deregulation of the labour market was implemented in 1991 with the Employment Contracts Act.[19] This Act reduced the role of unions drastically and enterprise bargaining replaced centralised bargaining. While the impact of reform is often difficult to establish due to the lack of a suitable counterfactual (see Gorter and Poot, 1999) we might expect that an increase in labour market flexibility would reduce the employment elasticity of pay to the extent that it measures non-competitive forces in the labour market. However, following arguments put forward by Blanchflower and Oswald (1994) and others, the elasticity is likely to be lower at high levels of unemployment. In the New Zealand case, the beginning and end points of the 1986-96 decade were relatively buoyant, while 1991 was characterised by severe recession (see also Table 2). Table 5 shows that the employment elasticity of pay is indeed lower in 1991 than in the other two years. Moreover, it can be argued that there is some evidence of the elasticity having been lowered by the introduction of labour market reform since the 1996 elasticity is lower than the 1986 one. However, it could be alternatively argued that elasticity increased between 1991 and 1996 due to the decentralisation of bargaining during this period, which allowed for more wage variation at the local level (see also Longhi et al. 2004 with respect to the case of Germany).

Given the proximate nature of the wage variable in our empirical work, as discussed in the previous section, it is useful to compare the employment elasticity of

[19] For a review of this reform and the subsequent reintroduction of multi-firm collective bargaining with the Employment Relations Act in 2000, see Morrison (2003).

the hourly wage with those of hours worked and annual income. Given equation (22), the elasticity with respect to hours is the difference of the elasticity with respect to annual income and hours worked. These elasticities are also reported in Table 5. All regressions are carried out with the specification of column (1) of Table 3. In general, hours worked do not respond strongly to variation in local employment rates, with the elasticity being significant in only about half the cases, and some of the significant coefficients are negative, suggesting a backward bending labour supply. However, the estimates of Equation (24) using hourly wages differ little from those using annual income.

Given that we have maintained a close link between the theoretical model (14) and the estimation equation (24), it is possible to calculate the values of the parameters of the quit-rate function that follow from the feasible GLS regression estimates. It can be easily derived from Equations (16), (17) and (18) that the elasticities of the quit-rate function (8) can be recovered as $\hat{\eta} = \hat{\alpha}\hat{\delta}$, $\hat{\nu} = \hat{\alpha}\hat{\xi}$ and $\hat{\mu} = \hat{\alpha}\hat{\varphi}$, where $\hat{\alpha} \equiv [2(1-\hat{\xi}-\hat{\varphi})]^{-1}$. Given that $\hat{\xi}$ = 0.045, $\hat{\varphi}$ = 0.113 and $\hat{\delta}$ = 1.894 in column (1) of Table 3, we find that $\hat{\nu}$ = 0.027 (0.039), $\hat{\mu}$ = 0.067 (0.017) and $\hat{\eta}$ = 1.124 (0.144), with standard errors (computed using the delta method) shown in parentheses. Local quits are, as expected, more responsive to local wage opportunities than to wages offered in other local labour markets. In turn, quit rates are more elastic with respect to wages offered elsewhere than with respect to a change in the local real non-wage income. The parameters $A_.$ and $B_.$ cannot be separately identified by regression equation (24).

Table 6 reports estimated quit-rate function parameters for all sub-groups derived from GLS estimation of (24) under inequality constraints (ξ, φ and δ were constrained to be greater than or equal to zero) Standard errors were computed with the delta method. It should be noted that, as the resulting quit-rate function parameters have been estimated indirectly from earnings outcomes and the spatial linkages between local labour markets, rather than directly labour turnover data, the estimates are likely to be rather imprecise. Nonetheless, it can be seen from Table 6 that the quit rate is responsive to the local wage and to the wage offered in other local labour markets and provides therefore some support for the theoretical model. The effect of non-wage income on quitting behaviour is generally not statistically significant. The difficulty in obtaining suitable data on this variable may have influenced the result.

TABLE 6 ABOUT HERE

5. Conclusion

In this paper we examined the role that local labour market conditions play in wage determination. We extended a labour turnover cost model proposed by Campbell and Orszag (1998). While their model is essentially non-spatial, in this paper we explicitly incorporated the interaction between local labour markets.

The theory was put to the test by means of a 'synthetic' micro-sample of 20,302 observations in 30 New Zealand local labour markets, derived from 1986, 1991 and 1996 New Zealand census data. The employment rate was found to enter the wage equation with a positive coefficient, consistent with the notion of a short-run wage curve. In addition, more remote labour market areas coincide with lower earnings, *ceteris paribus*. Moreover, local quits respond to available wages in other local labour markets.

A key conclusion of the study is that differences in the extent to which local labour market conditions affect the wages of subgroups of workers are important. Some evidence of a wage curve was found for men, but not for women, even after controlling for endogeneity of the employment rate. One explanation for this, consistent with the turnover cost model, is that the labour force participation of female workers is more elastic than that of males. A similar argument may also explain why the employment elasticity of pay is relatively low for workers aged under 26.

Furthermore, the earnings of non-Europeans, part-time workers and those with less education are more vulnerable to the state of the local business cycle. On the whole, our model and empirical evidence are consistent with the notion that local employers engage in monopsonistic competition with respect to the employment of less mobile workers. In general, the elasticities found for hourly earnings did not differ greatly from those found for annual income, suggesting that hours worked varied little with local labour market conditions.

Blanchflower and Oswald (1994) described their empirical findings on wages and unemployment as featuring 'two of the variables that most interest policymakers' (p. 1). The results of this study suggest that the wages of certain groups of workers are more responsive to changes in local employment prospects than those of others. Any attempt to reduce this vulnerability will then require increasing the geographical mobility of these workers. Future research needs to address the cause of the observed variation in the slope of the wage curve by estimating the quit-rate function directly using labour turnover data.

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Appendix *Definitions of the variables used*

Aged 26-40	Dummy variable identifying subgroups of workers who are aged 26-40.
Aged 41-60	Dummy variable identifying subgroups of workers who are aged 41-60.
d	Average travelling time T to all other urban areas, weighted by the working age population of the destination urban area.
European	Dummy variable identifying subgroups of workers who are European.
Full-time	Dummy variable identifying subgroups of workers who work 30 hours or more per week.
H	Mean weekly hours worked by persons in the subgroup. This refers to hours 'normally' worked per week in the 1996 Census. For 1986 and 1991, this refers to the predicted values from a regression of hours worked in the previous week (as supplied by the census question) on the normal 1996 hours for the corresponding subgroup.
I	Weighted average of the national growth rates over the previous year of employment in eight industry groups, namely, forestry and mining, manufacturing, electricity, building, wholesale and retail trade, transport, business services and personal services. The weights were calculated as the share of an urban area's employment in each industry in the particular year.
ι	Set of dummy variables for the census years, *i.e*. 1986, 1991 and 1996.
L_F	Proportion of working age females who are in the labour force in a given urban area and year.
L_M	Proportion of working age males who are in the labour force in a given urban area and year.
L_T	Proportion of working age people who are in the labour force in a given urban area and year.
Male	Dummy variable identifying subgroups of workers who are male.
N	Number of people aged between 15 and 59 in a given urban area and year, *i.e*. the working age population.
Occupation	Set of dummy variables for subgroups of workers in the following occupation groups: (1) Health professionals; (2) Teaching professionals; (3) Office clerks; (4) Salespersons, demonstrators and models; (5) Personal and productive service workers; (6) Skilled and semi-skilled production workers; (7) Construction workers; (8) Labourers and related elementary service workers.
P	National Consumers' Price Index value in a given year (base of 1 in 1986).
p	Average sale price of houses on the freehold open market for the half year ending June of the particular year, relative to the corresponding national value. Values refer to the territorial authority or authorities that match the relevant urban area most closely.
Qualified	Dummy variable identifying subgroups of workers with post secondary school qualifications.
r	Proportion of private dwellings in a given urban area and year that are rented.
T	Travel time T between two local labour market areas. Times are based on an average speed of 80 km h^{-1} but allow for extra time on congested roads and include a 3 hour ferry crossing where applicable.
u_F	Proportion of female labour force that is unemployed in a given urban area and year.
u_M	Proportion of male labour force that is unemployed in a given urban area and year.
u_T	Proportion of total labour force that is unemployed in a given urban area and year.
$\hat{w}$	Estimated real hourly wage for persons in the subgroup, calculated as $\hat{w} = \frac{Y}{(52H)P}$.
ω	Set of dummy variables for the following urban areas: (1) Whangarei; (2) Auckland; (3) Hamilton; (4) Tauranga; (5) Rotorua; (6) Gisborne; (7) Napier/Hastings; (8) New Plymouth; (9) Wanganui; (10) Palmerston North; (11) Wellington; (12) Nelson; (13) Christchurch; (14) Dunedin; (15) Invercargill; (16) Pukekohe; (17) Tokoroa; (18) Taupo; (19) Whakatane; (20) Hawera; (21) Feilding; (22) Levin; (23) Kapiti; (24) Masterton; (25) Blenheim; (26) Greymouth; (27) Ashburton; (28) Timaru; (29) Oamaru; (30) Gore.
W_R	Annual rainfall in millimetres for a given location and year.
W_S	Annual sunshine in hours for a given location and year.

Y Median income from all sources for persons in the subgroup over the past year.

Sources:
All variables were obtained from the 1981, 1986, 1991 and 1996 Censuses of Population and Dwellings, except for the following: *p*, which was taken from Quotable Value New Zealand's *Urban Property Sales Statistics*, W_R and W_S, which were supplied by the National Institute of Water and Atmospheric Research, and *I*, which was derived from both the census and Quarterly Employment Survey.

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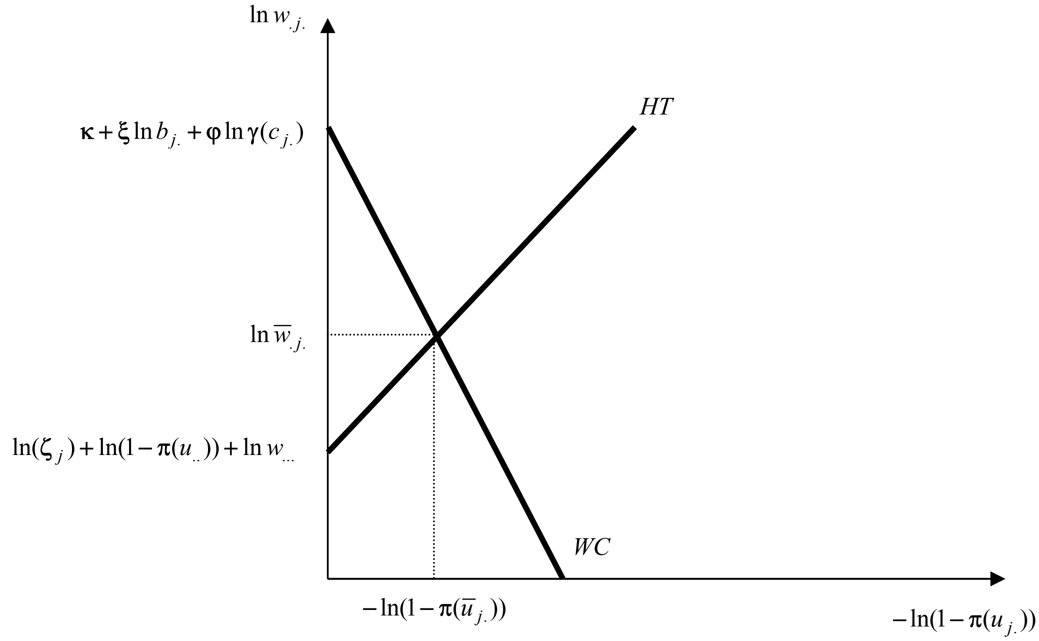

Figure 1
The wage curve (*WC*) and Harris-Todaro condition (*HT*)

Table 1a
Descriptive statistics: synthetic sample of 20,302 observations

Variable	1986		1991		1996	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
Y	14,070	6201	21,122	9747	21,582	11,072
H	36.56	10.53	36.07	11.49	34.75	13.48
w	7.378	2.856	7.434	2.568	7.175	2.887
Full-time	0.797	0.402	0.765	0.424	0.702	0.458
European	0.869	0.338	0.864	0.343	0.858	0.349
Male	0.386	0.487	0.370	0.483	0.367	0.482
Aged 26-40	0.334	0.472	0.352	0.478	0.340	0.474
Aged 41-60	0.311	0.463	0.356	0.479	0.355	0.479
Qualified	0.343	0.475	0.437	0.496	0.373	0.484
Health professionals	0.070	0.255	0.077	0.267	0.070	0.255
Teachers	0.079	0.270	0.095	0.293	0.096	0.295
Office clerks	0.371	0.483	0.363	0.481	0.332	0.471
Salespeople	0.092	0.289	0.105	0.306	0.124	0.330
Service workers	0.171	0.377	0.183	0.387	0.205	0.404
Production workers	0.136	0.343	0.097	0.296	0.093	0.290
Construction workers	0.035	0.184	0.032	0.175	0.020	0.139
Labourers	0.045	0.207	0.048	0.213	0.060	0.238

Table 1b
Descriptive statistics for the population of all wage and salary earners

Variable	1986		1991		1996	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
Y	15,652	6402	23,686	10209	25,355	12164
H	39.43	9.91	39.00	10.89	38.24	12.97
w	7.630	2.865	7.729	2.621	7.684	2.957
Full-time	0.857	0.350	0.830	0.376	0.780	0.414
European	0.868	0.338	0.872	0.335	0.871	0.335
Male	0.565	0.496	0.531	0.499	0.511	0.500
Aged 26-40	0.358	0.479	0.381	0.486	0.374	0.484
Aged 41-60	0.314	0.464	0.356	0.479	0.363	0.481
Qualified	0.380	0.486	0.475	0.499	0.410	0.492
Health professionals	0.039	0.195	0.045	0.208	0.041	0.198
Teachers	0.048	0.214	0.060	0.237	0.060	0.237
Office clerks	0.200	0.400	0.202	0.402	0.187	0.390
Salespeople	0.053	0.224	0.061	0.240	0.073	0.260
Service workers	0.101	0.302	0.110	0.313	0.125	0.331
Production workers	0.083	0.276	0.063	0.242	0.060	0.238
Construction workers	0.021	0.142	0.019	0.136	0.012	0.109
Labourers	0.028	0.165	0.030	0.172	0.038	0.191

Table 2
Descriptive statistics for the urban area level variables

Variable	1981		1986		1991		1996	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
u_T	0.043	0.010	0.070	0.011	0.143	0.016	0.083	0.013
u_M	0.041	0.010	0.054	0.009	0.135	0.018	0.078	0.013
u_F	0.046	0.011	0.092	0.016	0.153	0.018	0.088	0.015
d	8.813	2.800	8.508	2.955	8.415	2.970	8.334	2.963
p	1.041	0.150	1.087	0.238	1.124	0.275	1.066	0.351
r	–	–	0.244	0.032	0.245	0.034	0.275	0.028
W_R	–	–	1099.725	200.568	1117.388	261.5697	1221.870	325.380
W_S	–	–	2129.761	191.692	2008.319	205.181	2046.717	165.396
I	–	–	0.007	0.005	-0.016	0.003	0.032	0.002

Table 3
Results from estimation of the turnover cost model earnings equation

Regressor	Specification			
	(1) *GLS* Total	(2) *IV* Total	(3) *IV* Male	(4) *IV* Female
Constant	3.585 [c]	2.383	3.312 [b]	1.160 [a]
	(0.503)	(1.564)	(1.382)	(0.619)
$\ln(1-u)$	1.894 [c]	0.822	2.279 [b]	-0.743
	(0.181)	(1.396)	(1.145)	(0.573)
Aged 26-40	0.362 [c]	0.362 [c]	0.473 [c]	0.295 [c]
	(0.017)	(0.017)	(0.014)	(0.018)
Aged 41-60	0.399 [c]	0.400 [c]	0.532 [c]	0.327 [c]
	(0.020)	(0.020)	(0.019)	(0.019)
Qualified	0.132 [c]	0.132 [c]	0.133 [c]	0.114 [c]
	(0.004)	(0.003)	(0.005)	(0.003)
European	0.052 [c]	0.052 [c]	0.084 [c]	0.024 [a]
	(0.011)	(0.019)	(0.015)	(0.012)
Male	0.151 [c]	0.173 [c]	–	–
	(0.009)	(0.034)		
Full-time	-0.016 [a]	-0.016 [a]	-0.097 [c]	-0.018 [b]
	(0.009)	(0.009)	(0.025)	(0.007)
$\ln b \approx \ln p^{-1}$	0.045	-0.004	0.016	-0.035
	(0.037)	(0.072)	(0.063)	(0.024)
$\ln(\gamma(c)) \approx -d$	0.113 [c]	0.059	0.097	-0.019
	(0.030)	(0.076)	(0.066)	(0.033)
Number of observations	20,302	20,302	9784	10,518
$\overline{R}^2$	0.769	0.768	0.728	0.837
AIC	-3.296	-3.294	-2.968	-3.766
SC	-3.278	-3.275	-2.933	-3.733

Note: a, b and c refer to statistical significance at the 10%, 5% and 1% levels, respectively.
Robust standard errors are given in parentheses.
All equations also include occupation, time and urban area dummy variables.

Table 4
Correlations of instrumental variables

Instrument	Correlation	
	$\ln(1-u)$	ε
I	0.682 [c]	-0.003
W_S	0.246 [c]	-0.004
W_R	0.036 [c]	-0.003
r	0.166 [c]	0.001
$\ln(1-u)_{-1}$	0.038 [c]	0.005

Note: a, b and c refer to statistical significance at the 10%, 5% and 1% levels, respectively.

Table 5
Employment elasticities from GLS estimation
of the turnover cost model earnings equation for different groups

Group	Dependent variable			Number of observations
	$\ln \hat{w}$	$\ln H$	$\ln Y$	
Total sample	1.894 [c]	-0.073 [a]	1.821 [c]	20,302
1986	2.639 [c]	0.217 [c]	2.857 [c]	6379
1991	1.530 [c]	0.194 [c]	1.724 [c]	6762
1996	1.932 [c]	0.491 [c]	2.424 [c]	7161
Male	0.401 [a]	0.048	0.449 [a]	9784
Female	-0.072	-0.086 [b]	-0.158	10,518
Aged 15-25	1.382 [c]	0.311 [c]	1.608 [c]	6582
Aged 26-40	2.818 [c]	-0.204 [c]	2.353 [c]	6982
Aged 41-60	1.737 [c]	-0.163 [b]	1.384 [c]	6738
Low education	2.092 [c]	-0.043	1.855 [c]	10,790
High education	1.698 [c]	-0.038	1.454 [c]	9512
Part-time	3.053 [c]	-0.129	2.621 [c]	7842
Full-time	1.356 [c]	0.058	1.292 [c]	12,460
European	1.615 [c]	-0.064	1.342 [c]	12,757
Non-European	2.567 [c]	-0.053	2.474 [c]	7545
Health professionals	0.066	-0.101	-0.356	1950
Teachers	0.276	-0.031	0.220	2371
Office clerks	1.452 [c]	0.140 [b]	1.335 [c]	3103
Salespeople	-0.000	0.316 [c]	0.231	2714
Service workers	2.061 [c]	-0.342 [b]	1.601 [c]	3429
Production workers	1.452 [c]	-0.088	1.182 [c]	2997
Construction workers	0.583	-0.084	0.678	1220
Labourers	0.780 [c]	-0.189 [a]	0.653 [a]	2518

Note: a, b and c refer to statistical significance at the 10%, 5% and 1% levels, respectively.

Table 6
Quit-rate function elasticities from GLS estimation
of the turnover cost model earnings equation for different groups

Group	Parameter		
	η	ν	μ
Total sample	1.124 [a]	0.027	0.067 [a]
	(0.144)	(0.039)	(0.017)
1986	1.321 [a]	–	0.000
	(0.113)		(0.003)
1991	0.783 [a]	–	0.012 [a]
	(0.198)		(0.002)
1996	0.988 [a]	–	0.011 [a]
	(0.160)		(0.002)
Male	0.203 [b]	0.000	0.007
	(0.120)	(0.000)	(0.014)
Female	0.000	0.000	0.010
	(0.000)	(0.000)	(0.008)
Aged 15-25	0.706 [a]	0.000	0.007
	(0.117)	(0.000)	(0.014)
Aged 26-40	1.959 [a]	0.069	0.126 [a]
	(0.313)	(0.066)	(0.028)
Aged 41-60	1.026 [a]	0.012	0.079 [a]
	(0.145)	(0.039)	(0.021)
Low education	1.237 [a]	0.025	0.066 [a]
	(0.176)	(0.042)	(0.018)
High education	1.040 [a]	0.034	0.079 [a]
	(0.143)	(0.036)	(0.021)
Part-time	2.187 [a]	0.122	0.094 [a]
	(0.585)	(0.084)	(0.035)
Full-time	0.745 [a]	0.000	0.050 [a]
	(0.107)	(0.000)	(0.015)
European	0.955 [a]	0.019	0.072 [a]
	(0.123)	(0.034)	(0.017)
Non-European	1.693 [a]	0.073	0.086 [a]
	(0.330)	(0.063)	(0.036)
Health professionals	0.037	0.000	0.071 [a]
	(0.162)	(0.000)	(0.018)
Teachers	0.138	0.000	0.000
	(0.106)	(0.000)	(0.000)
Office clerks	0.901 [a]	0.042	0.079 [a]
	(0.212)	(0.059)	(0.021)
Salespeople	0.000	0.000	0.002
	(0.000)	(0.000)	(0.019)
Service workers	1.327 [a]	0.111 [b]	0.032
	(0.246)	(0.061)	(0.024)
Production workers	0.838 [a]	0.000	0.077 [a]
	(0.253)	(0.000)	(0.030)
Construction workers	0.292	0.000	0.000
	(0.180)	(0.000)	(0.000)
Labourers	0.393 [a]	0.003	0.000
	(0.155)	(0.056)	(0.000)

Note: Standard errors, calculated using the delta method, are given in parentheses.
a, b and c refer to statistical significance at the 10%, 5% and 1% levels, respectively.
Estimates of ξ, φ and δ used to calculate η, ν and μ were constrained to be greater than or equal to zero.
Estimates of ν cannot be obtained for separate years as ln$p$ is a location-invariant regressor.